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Corporate Financial Constraints, Minimum Wage Policies, and Employment

By *Alona Bilokha, Iftekhar Hasan, Stefano Manfredonia, Ronald W. Masulis*

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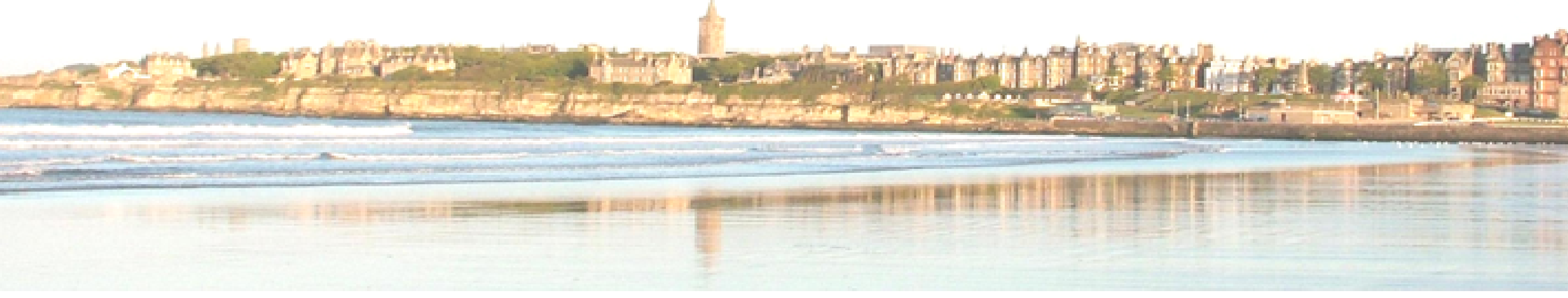

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Abstract

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Keywords: Local Labor Markets; Minimum Wage; Financial Frictions; Firm and Establishment Employment.
JEL Codes: G30, G32, J30, J38.

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1 Introduction

Minimum wage laws in the United States differ significantly from state to state, with the Federal government establishing a baseline minimum wage that states can choose to exceed. The introduction of The Raise the Wage Act in Congress on July 25, 2023, intensified minimum wage policy discussions by proposing to raise the federal minimum wage from \$7.25 an hour, first set in 2009, to \$17 an hour by 2029.[1] This congressional proposal was projected to directly raise the salaries of over 27 million U.S. workers, representing nearly 20% of the wage-earning workforce. As a result, minimum wages have grown in importance in current U.S. government policy debates (Zipperer, 2023).

One central point of contention in these debates involves the impact of minimum wage law changes on employment levels. An in-depth understanding of this relationship is of great importance to employers, lower-income workers, and policymakers, as it speaks directly to the welfare implications of minimum wage policies. Despite a large body of economic research investigating this relationship, a clear consensus has yet to emerge on whether minimum wage increases affect employment, with published evidence drawing fundamentally different conclusions.[2]

Our primary objective is to shed new light on this minimum wage-employment relationship by exploring the role of corporate financial characteristics at the individual firm level in determining firm employment responses to labor factor price shocks. In fact, the causal chain between minimum wages and employment studied in nearly all prior literature ignores the potential role that firms and their balance sheets can play

[1] For details, see `https://www.cbo.gov/publication/55681` (this reference was last accessed on December 2, 2025).

[2] For example, in a widely cited study, Card and Krueger (1994) analyze the effect of an increase in New Jersey's minimum wage and show that fast-food restaurants in this area increased employment by 13% relative to nearby Pennsylvania stores. Cengiz et al. (2019) reports no significant change in the total number of low-wage jobs during the five years after a minimum wage adjustment. Using survey information at the individual level, Clemens and Wither (2019) identify a negative effect on employment following the federal minimum wage increase during the 2007-2008 financial crisis. For reviews of the extensive literature on the effects of minimum wage policies on employment in the United States, see Neumark and Wascher (2008), Belman and Wolfson (2014), and Schmitt (2015).

in employment decisions. Yet, there are compelling reasons to believe that firm-level factors play a fundamental role in determining the impacts of minimum wage increases on aggregate labor demand. Firms decide whether to hire, fire, or retain employees and experience wide-ranging heterogeneity in their available resources, exposure to state regulations, and financial flexibility. From this perspective, recent empirical evidence emphasizes the importance of firms' financial constraints in understanding aggregate employment dynamics (Giroud and Mueller, 2017, 2019). A more precise understanding of the role of firm characteristics in this context can aid policymakers in designing more effective minimum wage policies and help to reconcile previously mixed empirical findings in the literature. Our paper aims to answer this research question by analyzing whether corporate financial constraints are essential in explaining employment dynamics following minimum wage increases and uncovering the underlying mechanisms that link the two.

Although previous research in finance has predominantly examined how restrictions on accessing external finance affect firms' capital investment decisions (e.g., Hubbard, 1998), financial constraints can also distort other major firm decisions, particularly those with cash flow implications (Almeida et al., 2024). Specifically, a decline in firm-level employment can follow a minimum wage hike if firms are unable to cover these additional labor expenses solely from price hikes or internal resources. This phenomenon specifically arises when firms face serious financial constraints that prevent them from borrowing against their future cash flows. More importantly, even when corporations have internal resources available, financially constrained firms are expected to place a relatively high premium on preserving liquidity (Almeida et al., 2004; Faulkender and Wang, 2006; Denis and Sibilkov, 2010; Dasgupta et al., 2019).

To advance our research agenda, we created a new database that contains information on employees at the establishment level as well as corporate financial characteristics. This approach is more precise than solely examining firms and their headquarters location, as minimum wage laws are based on their employees' locations, and many firms have employees spread over multiple states and internationally. For this analysis, we extract information on all domestic establishments that belong to publicly traded U.S. firms, as well as data on firm balance sheet characteristics.[3]

Using this database, we explore two alternative natural experiments. First, we adopt a unique quasi-experimental setting to provide causal evidence on the relationship between minimum wage changes, financial constraints, and employment levels. We examine the increase in the federal minimum wage during the 2007-2008 financial crisis and compare the employment dynamics of establishments located in states affected by the federal minimum wage change to establishments located in states unaffected by the federal wage increase (i.e., states where a state's minimum wage was equal to or exceeded the new minimum wage level mandated by the federal government) (Gustafson and Kotter, 2023), while controlling for aggregate firm dynamics. In fact, during the financial crisis, the establishments in affected states experienced significantly more growth in their *effective* average wage rates compared to the establishments in unaffected states (Clemens and Wither, 2019).

The primary advantage of this setting is that it also provides a plausibly exogenous shock to firms' access to external finance, allowing us to examine how an exogenous reduction in the distance from financial constraint affects the responses of employment to changes in the minimum wage. More specifically, we follow the previous literature (Almeida et al., 2011; Benmelech et al., 2019; Duval et al., 2020) and exploit the heterogeneity across firms in the relative portion of their long-term debt that matures at the onset of the financial crisis. The rationale behind this approach is that firms with large amounts of debt maturing during this financial crisis period are generally unable to roll over their maturing debt given the serious disruption to capital markets,

[3] While several existing studies use establishment-level databases to examine the effects of minimum wage policies on employment and wages (e.g., Chava et al., 2023; Bossler and Gerner, 2020; Gopalan et al., 2021; Dustmann et al., 2022), our paper is the first to merge establishment-level data with key details from a firm's balance sheet. This integration of firm- and establishment-level data enables us to obtain new insights into the interplay between minimum wage policies and corporate employment dynamics, explicitly shedding light on the pivotal role of corporate financial constraints in this context. Furthermore, corporate balance sheets are utilized in our analysis not only to examine the heterogeneous effects of minimum wage policies on employment, but also to illustrate how exposure to these policies influences corporate cash holdings and financing decisions.

causing these firms to experience more binding financial constraints. Thus, these firms are compelled to modify their behavior to a greater extent than similar firms without the need to refinance their long-term obligations during the crisis period.

Using this experimental setting, we discover that increasing the minimum wage during this critical period significantly reduces employment in treated establishments. We observe a 1.4% decrease in employment in establishments located in states affected by the new federal minimum wage. Furthermore, we show that this effect is immediate and persists for several years. More importantly, we find that establishments belonging to firms with a large amount of debt maturing at the onset of the financial crisis and located in states affected by the new federal minimum wage requirements experienced greater declines in their establishment employment levels. Specifically, our analysis shows that an increase of one standard deviation in a firm's financial constraint measure leads to a 1.1% decline in its treated establishment' employment levels. This result is driven by establishments with low cash holdings at the onset of the crisis, consistent with the hypothesis that financially constrained firms seek to preserve their cash reserves after a minimum wage hike. In addition, a counterfactual exercise provides additional evidence that minimum wage policies could substantially negatively affect *aggregate* employment in the presence of widespread financial constraints.

A potential issue with this quasi-natural experiment is that, although the debt maturity structure is predetermined at the onset of the crisis, it may not be exogenous to other firm characteristics that influence employment responses to minimum wage policies. More specifically, firms with high levels of short-term debt could systematically differ from other firms in ways that affect employment decisions but are unrelated to their financial constraints, such as a strategic focus on short-term performance. However, we demonstrate that constrained and unconstrained firms, defined by their debt maturity structure, are similar in several observable characteristics before the onset of the financial crisis. Furthermore, our analysis of the triple interaction terms, defined by yearly indicators, the measure of firm financial constraints, and exposure to the federal minimum wage increase, supports the parallel trends assumption, further validating our approach.

Despite the strong internal validity of this setting, this investigation focuses on a historical period marked by unusually high financial constraints. As a complementary setting, we then examine a more typical and broadly generalizable environment in which firms are less financially constrained, allowing us to better capture the behavior of unconstrained firms and assess the external validity of our findings. More specifically, we exploit staggered changes in state minimum wages across states and standard measures of financial constraints. Extending the previous literature, our combined firm- and establishment-level database enables us to compare similar establishments belonging to the same firm, but located in different states, before and after a state minimum wage change, while controlling for aggregate firm-level trends.[4]

Using the above empirical strategy, we show that a rise in the minimum wage does not affect establishments' employment levels in subsequent years in normal economic times. However, consistent with the results in the financial crisis setting, we document a significant adverse effect on the employment levels of constrained firms following minimum wage rises. In economic terms, our findings indicate that an increase of one standard deviation in a firm's financial constraint, represented by alternative conventional financial constraint metrics, leads to an average 0.9% reduction in employment at the affected establishments following a one-dollar increase in the minimum wage.

While examining the marginal effects of minimum wage rates on employment across the entire range of individual firm financial constraints, we find some positive and significant effects on employment at unconstrained firms in normal economic times. We provide evidence that this outcome is unlikely to be entirely explained by the reallocation of minimum wage workers from constrained establishments to unconstrained ones using aggregate county-level regressions. Interestingly, we show that the rise in

[4] While this approach addresses several endogeneity issues, it does not definitively establish causality due to potential correlations between firm financial constraints and other firm and establishment characteristics that may influence employment decisions. For example, financially constrained and unconstrained firms can differ along multiple dimensions, including labor adjustment costs, managerial quality, and risk preferences, which could independently influence employment levels after a minimum wage change.

employment is concentrated among establishments of unconstrained firms located in areas with a larger supply of potential minimum-wage workers, as well as in areas with a higher hiring gap and elevated quit-to-hire ratios. This finding is consistent with the theoretical predictions of well-established job search models, which argue that a minimum wage increase can lead to higher overall employment by firms motivated to post more vacancies, thereby allowing them to fill more positions, as unemployed workers adjust their search behavior in response to higher wages (Burdett and Mortensen, 1998; Flinn, 2006; Giuliano, 2013; Manning, 2021). Similarly, this evidence aligns with job search models that incorporate frictions, suggesting that minimum wage policies can increase employment by reducing labor turnover and improving matching efficiency (Portugal and Cardoso, 2006; Brochu and Green, 2013; Dube et al., 2016, 2019). In contrast, we find no evidence that establishments of unconstrained firms increase employment after a minimum wage hike in monopsonistic labor markets, where employers exert significant wage-setting power, as recently documented by Azar et al. (2024).

A firm-level analysis offers further insight into the mechanisms by which minimum wage policies influence the employment dynamics of financially constrained firms. We find that liquidity buffers decline following minimum wage increases. Since these policies reduce firms' available internal funds, affected firms also turn to external financing to meet their funding needs and build a financial buffer to manage future wage pressures and absorb higher operating costs. Consistent with this hypothesis, we document an increase in the use of long-term debt after increased exposure to minimum wage policies. Because financially constrained firms tend to preserve liquidity and reduce employment, while unconstrained firms expand it during normal economic times, these adjustments appear to be primarily driven by the latter sample of firms.

We acknowledge several inherent limitations in our empirical exercise. Although our two quasi-experimental settings are designed to mitigate endogeneity concerns, neither can fully rule out the influence of unobserved firm- or local-level shocks correlated with both minimum wage exposure and corporate financial constraints. In addition, no single proxy can perfectly capture the multidimensional nature of corporate financial constraints. We address these issues through complementary identification strategies, alternative measures of financial constraints, and an extensive battery of robustness checks, but readers should interpret our findings with these limitations in mind.

2 Previous literature and contribution

Our paper contributes to several distinct strands of literature. In finance, a growing body of work has analyzed the effect of minimum wage laws on corporate policies. Gustafson and Kotter (2023) find that increases in minimum wages in the United States lead public firms to cut capital expenditures. Geng et al. (2022) find opposite results by analyzing the effect of minimum wage increases in China and show for a manufacturing firm sample that a rise in the minimum wage is associated with increased capital investment and innovation. Using a similar setting, Hau et al. (2020) report that minimum wage increases in China accelerate input substitution from labor to capital, reduce employment growth, and accelerate total factor productivity growth. They also show that this effect is particularly strong among less productive firms under private Chinese or foreign ownership, but these patterns do not occur among state-owned enterprises. Examining a large and persistent minimum wage increase in Hungary, Harasztosi and Lindner (2019) show that firms responded by substituting capital for labor. Using a comprehensive data set from the hotel industry, Agarwal et al. (2024) exploited staggered state-level changes in minimum wages in the United States from 2000 to 2008 to find that doubling the minimum wage reduces average hotel revenues by 6% per year and occupancy rates by 3.1%. Chava et al. (2023) show that increases in the federal minimum wage negatively affect the financial health of small, predominantly private businesses in the affected states. Importantly, their findings also indicate a causal relationship between minimum wage hikes and increased small business closure rates.

The only study that takes into account the financial constraint of the firms found in this literature is Arabzadeh et al. (2024), who analyzes a minimum wage change in Germany and uses employee-employer information and a structural model to investigate the relationship between minimum wage changes, financial frictions, and within-firm wage dispersion. They find that within-firm wage dispersion declines more with higher minimum wages when firms are financially constrained. However, other studies provide suggestive evidence that financially distressed and less productive firms are more negatively affected by minimum wage policies. For example, examining Portuguese firms, Alexandre et al. (2022) show that minimum wage hikes reduce employment growth and profitability, particularly among financially distressed firms. Analyzing German companies, Dustmann et al. (2022) document that following minimum wage increases, small and low-productivity establishments experience lower employment and a higher exit rate. Studying Israeli firms, Drucker et al. (2021) find that firms owned by lower-income individuals face lower profitability after a minimum wage increase. Rao and Risch (2025) provide evidence that an increase in minimum wage causes firm exits. They also show that these exits are concentrated among the least productive small firms.

Our paper provides a comprehensive analysis of how reduced access to external financial resources affects employment dynamics following minimum wage changes. Using a unique combination of establishment- and firm-level data on U.S. public companies, we provide evidence that minimum wage policies have negative average effects on employment during crisis periods in capital markets, when access to external financial resources can be seriously restricted. By exploiting firm debt maturity structures at the onset of the crisis, we establish a causal relationship between corporate access to external financing and employment changes following a minimum wage increase. Furthermore, we provide evidence that the precautionary savings motives of financially constrained firms helps explain these patterns.

As a complementary analysis, we exploit staggered changes in minimum wage policies across U.S. states and find no average effect on employment during normal economic times. However, we show that this average effect masks substantial heterogeneity among firms: consistent with our earlier findings, establishments belonging to financially constrained firms experience a significant decline in employment, while establishments of unconstrained firms increase employment levels. This latter result offers an opportunity to explore and test alternative economic arguments from the minimum wage literature that suggest that minimum wage policies can lead to increased employment and reconcile some of the existing conflicting theories and evidence. In addition, a firm-level analysis offers new insights into how exposure to minimum wage policies affects a firm's financial performance. Among other findings, we demonstrate that minimum wage increases place pressure on the firm's internal resources, compelling these firms to seek additional external capital.

Our paper contributes to the literature that analyzes the impact of financial frictions on real outcomes. Duval et al. (2020) show that firm financial constraints during the Great Recession crisis produced a persistent negative effect on firm productivity and innovation outcomes. Benmelech et al. (2019) provide evidence that the lack of access to credit combined with financial frictions affected firm employment during the Great Depression. Caggese et al. (2019) use matched employer-employee data from Sweden and find that financing constraints push firms to sub-optimally fire short-tenured workers who can offer high expected future productivity. Chodorow-Reich (2014) shows that firms with weaker lender relationships had more difficulty obtaining loans and experienced higher interest rates after the Lehman bankruptcy shock. They also exhibited greater reductions in employment compared to companies with stronger lending relationships. Giroud and Mueller (2017) discovered that businesses belonging to highly leveraged firms suffered significantly greater reductions in employment when faced with a decrease in local consumer demand. Gilchrist et al. (2017) document that firms with liquidity constraints raised prices during the 2008 financial crisis, while their unconstrained counterparts lowered prices. In a recent related paper, Almeida et al. (2024) find that funding frictions can limit firms' short-term investments in receivables and inventories, reducing their production capacity. Furthermore, financial constraints have been shown to have a significant impact on various firm decisions, including choices about investment and capital structure, as well as on stock returns (Hennessy and Whited, 2007; Lamont et al., 2001; Cao et al., 2019).

Our analysis reveals a previously unexplored setting where financial frictions can produce real impacts. More specifically, we analyze a factor price shock to corporate internal resources that forces financially constrained firms to reduce their employee headcount to preserve liquidity. By doing so, we provide evidence that borrower financial weaknesses and the inability to access financial markets provide a plausible explanation for the rise in unemployment after a minimum wage hike (Clemens and Wither, 2019). This result is particularly relevant, considering the challenges that a large body of literature face in explaining the mixed results obtained when analyzing minimum wage changes over different countries and time periods (Belman and Wolfson, 2014).

3 Conceptual Framework

We argue that corporate financial constraints play a central role in shaping how minimum wage policies affect employment. An increase in the minimum wage, *ceteris paribus*, raises labor costs and requires firms to finance a higher wage bill. Although the classical literature emphasizes the impact of financial constraints on investment decisions, recent studies show that financial frictions also distort other key corporate choices, particularly those with direct cash-flow implications (Almeida et al., 2024).

Financially constrained firms place a high premium on preserving liquidity (Almeida et al., 2004; Faulkender and Wang, 2006; Denis and Sibilkov, 2010; Dasgupta et al., 2019). Beyond facing limited access to external capital, these firms are often reluctant to draw down internal cash reserves, viewing them as crucial buffers against future shocks and as safeguards for operational flexibility. This behavior is consistent with the literature that emphasizes the precautionary value of cash holdings: constrained firms preserve cash to mitigate disruptions when credit markets tighten (Campello et al., 2011), protect investment plans against adverse cash-flow conditions (Almeida et al., 2014), and ensure the availability of resources in uncertain economic environments (Nikolov et al., 2019; Lins et al., 2010; Rochet and Villeneuve, 2011).

When faced with higher wage costs, firms can, in principle, adjust their liquidity levels along two margins: by increasing prices or by adjusting employment. Although some studies document a partial price pass-through mechanism following minimum wage increases (Fougère et al., 2010; Leung, 2021; Rao and Risch, 2025), this adjustment is generally incomplete, especially in competitive or regulated product markets, and during economic downturns, when weak demand and tighter credit conditions limit firms' ability to pass higher labor costs onto consumers. In addition, in some industries, a substantial time lag between production and sales further amplifies concerns that higher revenues may be insufficient to cover rising labor expenses (Barrot and Nanda, 2020). As a result, higher wages translate into tighter liquidity conditions rather than being fully absorbed through prices.

In this environment, financial constraints become pivotal. For financially constrained firms, higher labor costs immediately tighten operating conditions by reducing internal funds available to support ongoing operations and to buffer against adverse, unexpected events. Because these firms face a high marginal value of cash holdings and an elevated shadow cost of approaching a binding financial constraint, absorbing wage increases through internal resources can be particularly costly. As a result, preserving liquidity can take precedence over other objectives, and thus, constrained firms are expected to adjust along the employment margin more aggressively than unconstrained firms, for whom this shadow cost is negligible.[5]

For financially unconstrained firms, the mechanism operates differently. Unconstrained firms can finance higher wage bills through internal liquidity or tapping external credit without facing a meaningful increase in the marginal cost of funds, since they remain far from any binding financial constraint. Still, in standard models, higher minimum wages increase labor costs and should reduce employment (Stigler, 1946). However, empirical evidence often finds muted average employment effects. This pattern can be reconciled by noting that unconstrained firms are better positioned to offset higher wages through channels such as reduced employee turnover, improved

[5] Constrained firms may also seek to avoid costly external financing, especially when capital markets are stressed or when firm-level financial conditions have deteriorated.

worker productivity, and exploiting an expanded pool of job seekers (Portugal and Cardoso, 2006; Brochu and Green, 2013; Dube et al., 2016, 2019; Coviello et al., 2022; Ku, 2022). These mechanisms can partially or fully counteract the direct cost effect of higher wages.

Together, this framework predicts heterogeneous employment responses to minimum wage policies. While higher minimum wages raise labor costs for all firms, financially constrained firms face an additional effective cost due to the elevated shadow cost of approaching a binding financial constraint. As a result, they are expected to cut employment more aggressively than unconstrained firms. In contrast, financially unconstrained firms remain far from any binding constraint and can finance higher wage bills using internal liquidity or external credit, since it does not materially raise their marginal cost of funds. These firms may reduce employment to a lesser extent, keep employment unchanged, or even expand employment when increased wages attract additional job seekers and reduce turnover. A simple theoretical model formalizing these mechanisms is provided in Online Appendix A.

4 Data

For our analysis, we collect information on minimum wage policies across individual U.S. states, listed firms' business establishment characteristics by state, and the balance sheet characteristics of the listed firms to which these establishments belong. Our variables and their data sources are described in more detail in the following.

Minimum wage policies. We obtain information on minimum wage policies from Vaghul and Zipperer (2021), who document state-level variations in minimum wage laws from 1974 to 2020. In our first quasi-experimental setting, we use this database to identify the states affected by the federal minimum wage increase during the financial crisis, that is, states whose effective minimum wage rate was lower than the new federal minimum wage. In our second quasi-experimental setting, we use information on yearly state-level minimum wage rates from 1990 to 2020, as establishment-level data are available only for this period.[6]

Counties along the state border. Minimum wage policies can be endogenous in some dimensions, since they are strongly tied to local economic conditions. To address this issue, in our preferred specifications, we follow the previous literature and focus our attention on geographically adjacent counties located along states' borders to ensure that the omitted local economic variables do not affect our results (Card and Krueger, 1994; Huang, 2008). In this way, we compare counties that are heterogeneously exposed to minimum wage treatments while they are geographically located in adjacent states, where we expect both observable and unobservable local economic conditions to be similar.

We also take into account the concern raised in Dube et al. (2016) that counties on state borders in the western U.S. are much larger and irregular in shape, and as such, they may not always share the same local economic conditions. To address this concern, Dube et al. (2016) further investigates counties on opposite sides of a state border by requiring their centroids to be within 75 km of each other. This 75 km distance cutoff is determined through a data-driven randomization inference procedure, which minimizes the mean squared error of the estimator. We highlight these counties by shading them in dark green in Figure OA B1.

Establishment level information. We gather data on the establishments of public corporations in individual states from the National Establishment Time Series (NETS) database (Addoum et al., 2020, 2023; Barkai and Panageas, 2025). It contains comprehensive information on the characteristics of publicly and privately owned establishments, including their locations from 1990 to 2020. With these data, we can accurately analyze the impact of minimum wage changes on employment levels.

One advantage of this database is that it is not subject to survivorship bias, which

[6] The exclusion of the 1980s is unlikely to materially affect our results, as there were minimal changes in state minimum wage rates before 1990 (Neumark et al., 2014; Allegretto et al., 2017).

is a major consideration in our analysis of the effect of minimum wage policies on the number of operating establishments. On the other hand, while the NETS database provides broad coverage and detailed establishment-level information, it also presents some limitations. In particular, employment data are partly based on model-based estimates rather than directly reported figures, especially for smaller or single-unit establishments. As discussed in Barnatchez et al. (2017), these estimates can introduce significant measurement error, particularly in time series analysis. However, our focus on publicly traded firms, whose employment and location information are more accurately tracked, mitigates these concerns. Furthermore, we conduct our empirical analysis using multiple data sources and aggregation levels.

Corporate balance sheet characteristics. From the Compustat database, we obtain comprehensive information regarding the total number of firm employees and balance sheet characteristics of publicly traded U.S. companies. With these data, we develop alternative metrics to evaluate financial frictions at the corporate level.

For our first quasi-experimental setting, we construct an exogenous measure of financial frictions based on ex-ante variation across firms regarding their long-term debt maturing during the Great Recession financial crisis period. The rationale behind this approach is that this financial crisis is unexpected. Thus, firms cannot deliberately schedule their debt to mature just before the onset of the crisis to avoid worsening credit conditions when the risk of debt rollover suddenly increases. The debt structure of firms before this unexpected event is also unlikely to be correlated with other unobserved firm characteristics or with each establishment's exposure to changes in minimum wages (Duval et al., 2020; Almeida et al., 2011; Benmelech et al., 2019). We investigate this claim in more detail in our empirical analysis.[7]

In our second quasi-experimental setting, we propose three different variables to

[7] Figure OA B2 shows the long-term debt maturity distribution of our sample firms. As expected, there is substantial heterogeneity in maturing debt levels over time. We exploit this heterogeneity in our empirical analysis by analyzing the amount of debt maturing during the Great Recession period versus in the other years.

measure corporate financial constraints that are widely used in the finance literature. Our first metric uses firm size as an indicator of financial friction, as previous research reports that smaller firms are more financially constrained (Gertler and Gilchrist, 1994; Siemer, 2019). We also use the Whited and Wu (WW) financial constraint index (Whited and Wu, 2006). The authors propose estimating an Euler equation derived from a structural investment model to create this measure. The index is constructed from six components: cash flows, assets, dividends, debt, and sales growth measured at industry and firm levels. As a third financial constraint measure, we include the Hadlock and Pierce size-age (SA) index (Hadlock and Pierce, 2010). The SA index is constructed by sorting firms according to firm characteristics closely linked to financial constraints. Hadlock and Pierce (2010) identified firm size and age as the factors most strongly associated with financial constraints. These two characteristics are much less likely to be endogenous than other variables that are commonly used to estimate firms' financial constraints, such as cash levels and leverage, which can be subject to discretionary decisions made by a firm's management and other limitations (Hoberg and Phillips, 2016; Buehlmaier and Whited, 2018). The SA index suggests that financial constraints decrease substantially as young and small firms mature and grow.

Although these measures are widely used in the corporate finance literature, we acknowledge that no single proxy can perfectly capture the underlying degree of financial constraint of a firm. Each metric reflects a particular aspect of financing frictions, such as firm size, age, or access to external markets, and may be sensitive to firm or time-specific factors. To mitigate these concerns, we employ additional complementary indicators and verify that our results remain qualitatively consistent across all specifications. This approach ensures that our findings do not depend on any single definition of financial constraints. More specifically, we consider older or less widely used measures of corporate financial constraints, such as the Kaplan-Zingales (KZ) index (Kaplan and Zingales, 1997), a composite index of various financial constraint indexes, a measure of financial constraints based on a textual analysis of 10-K filings (Bodnaruk et al., 2015), and indicators of whether a firm has issued a bond (Faulkender and Petersen,

2006) or obtained a syndicated loan (Campello et al., 2010; Acharya et al., 2013).

Finally, we collect additional firm variables from Compustat to better understand how exposure to minimum wage changes affects overall corporate policies and performance, including liquidity holdings, financial leverage, capital investment, and research and development expenditures.

Summary statistics. We use these data sources to understand how exposure to minimum wage regulation affects corporate employment decisions. To do so, we merge the NETS database with minimum wage levels and information on corporate balance sheet characteristics from Compustat.[8] We remove establishments that belong to firms that operate in the highly regulated utility and financial sectors (*SIC* code equal to 49 and 60), and winsorize all the variables related to establishment and firm characteristics at the 1st and 99th percentiles. Our final database comprises 2,340,503 establishment-year observations, 231,552 establishments, and 5,615 firms. We analyze these data over the period spanning 1990 to 2020. We report summary statistics for this sample in Panel A of Table OA B1. A description of the variables is reported in Panel A of Table OA B2.[9]

We use the establishment-level database to analyze whether establishments exposed to increases in minimum wages change their employment levels and whether corporate balance sheet characteristics explain these employment dynamics. We also aggregated the data to the firm level to further evaluate our findings and gain insight into how exposure to minimum wage adjustments impacts firm performance and policies. This database comprises 54,829 firm-year observations that span the period 1990-2020. We

[8] We use the legal business names of the establishments for cross-referencing with the Compustat database. A potential limitation of this approach is that we do not match subsidiaries with names different from those of their parent companies. However, subsidiaries often face different financial constraints than their parent companies, and many have significant operational autonomy. Also, we cannot merge the observations of companies that have changed their names over time. Nevertheless, our database remains substantially larger than the establishment coverage found in the U.S. Census Bureau's Longitudinal Business Database (LBD), which is widely used in the finance and economics literature (e.g., Giroud and Mueller, 2017).

[9] Additional details about the construction of our final datasets are found in Online Appendix C.

winsorize the corporate-level variables at the 1st and 99th percentiles and report the summary statistics in Panel B of Table OA B1. A detailed description of the variables included in our sample is reported in Panel B of Table OA B2.

5 The financial crisis and the federal minimum wage

We focus on a unique experiment that allows us to understand whether financial constraints *causally* affect establishment employment. For this purpose, we exploit two sources of exogenous variation in our sample period. One comes from an increase in the federal minimum wage during the Great Recession, which heterogeneously affected establishments across individual states depending on whether state minimum wage rates are below the new federal minimum wage requirement. This crisis period is particularly relevant for our analysis, as it is characterized by severe information asymmetries, uncertainty, and tight credit market conditions (Brunnermeier, 2009; Bernanke, 2023). The second source of exogenous variation comes from individual corporate debt structures and the ex-ante variation in a firm's long-term debt maturing in the crisis period. Since the financial crisis hit the U.S. economy unexpectedly, it meant that managers could not preemptively adjust their corporate debt structure to reduce their debt rollover risk during the financial crisis period.

For this exercise, we limit our sample period to focus on the 2007-2008 financial crisis period and analyze the broader 2003-2011 period. We report the summary statistics for the main variables used for this exercise in Table OA B3.

5.1 Federal minimum wage increase

The U.S. federal minimum wage change implemented during the Great Recession. Before the financial crisis, the federal minimum wage was set at $5.15 per hour, a rate established in 1997. This rate remained unchanged until 2007. Figure 1(a) illustrates the evolution of the federal minimum wage over the years of our sample period.

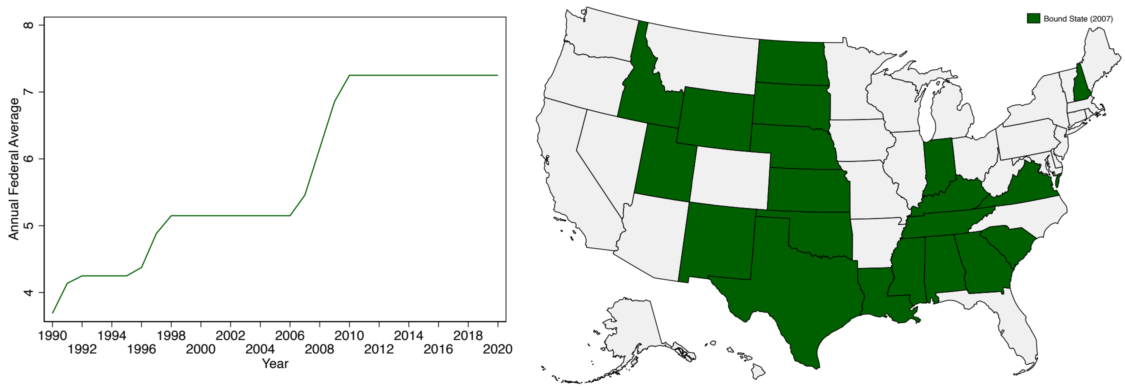
At the time of the federal minimum wage change, which occurred at the onset of the financial crisis, some states had minimum wage rates higher than the new federal minimum wage. As a result, the effective minimum wage in these states remained unaffected by the federal minimum wage increase. Figure 1(b) illustrates the states affected by this increase in the federal minimum wage.

The rationale behind our identification strategy is straightforward. After the minimum wage increase, firms may want to adjust their establishment employment levels due to the relative changes in local labor costs. To identify the impact of the federal minimum wage increase, we compare similar establishments in states affected by the wage increase before and after its implementation to establishments in states unaffected by the federal regulatory change (Clemens and Wither, 2019; Gustafson and Kotter, 2023).

As we see in Figure OA B3(a), after the onset of the financial crisis, the establishments of the affected states experience a greater growth in their *effective* minimum wage rates compared to those of the non-affected states, consistent with the fact that the regulatory shock has a tangible effect. More specifically, in the month before the federal minimum wage change, the average effective minimum wage rate was \$5.16 in affected states versus \$6.81 in unaffected states (a difference of \$1.65). At the end of the financial crisis, the effective minimum wage in the affected states was \$6.82 compared to \$7.37 in the unaffected states (a difference of \$0.55). Importantly, this convergence of more than \$1 was not due to any minimum wage change initiated by individual states, but is simply the result of the change in the federal minimum wage. We plot these divergent patterns in Figure OA B3(b), where we demean each time series using the pre-crisis average effective minimum wage rates for the treatment and control groups.

Employment dynamics after the federal minimum wage rise. To determine whether employment levels of establishments in affected states decrease, we estimate the following Equation:

Figure 1: The federal minimum wage change



(a) Changes in the federal MW

(b) Treated vs untreated states

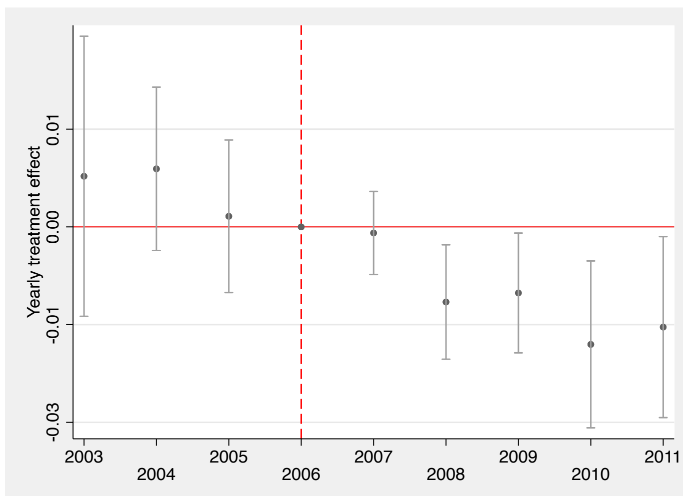
Notes: Figure 1(a) shows changes in the federal minimum wage over our sample period. Figure 1(b) illustrates the states that fall into the categories of affected (treated) and unaffected (untreated) in our empirical setting.

$$Log(Employment)_{i,t} = \sum_{t=2003, t\neq 2006}^{t=2011} \beta_t Treated_s \times Year_t + \eta_i + \theta_{f,t} + \epsilon_{i,t} \quad (1)$$

In Equation (1), *Log(Employment)* is the natural logarithm of the number of employees in establishment i at time t. *Treated* is an indicator variable that equals one if the establishment is located within a state treated by the rise in the federal minimum wage in 2007 and is zero otherwise. In this model, we interact this variable with a complete set of year-fixed effects using the year before the change in the federal minimum wage as the reference year. η_i and $\theta_{t,f}$ are respectively establishment and firm times year fixed effects. These controls allow us to compare the same establishment before and after an increase in the minimum wage, while controlling for aggregate firm dynamics. We further focus on counties located along state borders, which are similar in terms of economic characteristics. Therefore, the coefficients β_t report the differential effects of belonging to a state affected by the federal minimum wage increase on establishment employment for each particular year compared to the year before the change in the federal minimum wage.

We present the β coefficients estimated from Equation (1) and the relative confidence intervals in Figure 2. The graphs provide several striking findings. First, yearly point estimates show significant negative effects every year after the rise in the minimum wage. In terms of economic magnitude, we find that establishments in a state affected by a rise in the federal minimum wage decrease their yearly employment level by 0.3 % of their average employment level. Also, we find that the yearly treatment effects are not significant before the federal minimum wage change, suggesting that the parallel trend assumption holds and that we are comparing economically similar establishments.

Figure 2: Yearly treatment coefficients - Financial crisis



Notes: Figure 2 shows the yearly treatment effects from Equation (1). The outcome variable is the natural logarithm of employment in establishments. The treatment variable is an indicator variable that equals one if the establishment is located within a state affected by the rise in the federal minimum wage in 2007 and is zero otherwise. The regression includes establishment and firm times year fixed effects. We focus on counties located along state borders. The plot exhibits yearly point estimates and shows 95% confidence intervals based on standard errors clustered by state.

Since we show the effect is long-lasting and persistent, we consider a more aggregated difference-in-differences approach and estimate Equation (2) reported below:

$$Log(Employment)_{i,t} = \beta Treated_s \times Post_t + \eta_i + \theta_{f,t} + \epsilon_{i,t} \tag{2}$$

In this setting, the variable $Post$ is an indicator equal to 1 for 2007 and all subsequent years, and 0 otherwise. We report the estimates for the full sample, the sample of counties along state borders, and county pairs located within 75 km of each other in Table OA B4. In line with the results of the earlier event study, we find that establishments in states affected by a rise in the federal minimum wage decreased their employment levels. More specifically, according to the coefficient estimate reported in the last column, employment decreases by 1.4 %.

5.2 Financial constraints during the financial crisis

Based on our theoretical framework, we expect the negative effect we document to be more pronounced in establishments belonging to financially constrained firms. We test the role of this firm-level factor in this experimental setting by exploiting an exogenous measure of a firm's ability to access external capital, namely the ex-ante variation in long-term debt levels that mature at the onset of the financial crisis. This measure is expected to be unrelated to corporate investment prospects and other corporate characteristics, while affecting a firm's need for credit intermediation. In fact, firms with large amounts of debt maturing during this financial crisis period are generally unable to roll over their maturing debt, given the serious disruption to capital markets around the crisis period, causing these firms to experience more binding financial constraints. Thus, these firms are compelled to modify their behavior to a greater extent than similar firms without the need to refinance their long-term obligations during the crisis period (Almeida et al., 2011; Benmelech et al., 2019; Duval et al., 2020).

We use this exogenous measure of corporate financial frictions and estimate the following Equation:

$$\begin{aligned} Log(Employment)_{i,t} = {} & \beta_0 Treated_i \times Post_t \\ & + \beta_1 Treated_i \times Post_t \times Constraints_f + \eta_i + \theta_{f,t} + \epsilon_{i,t} \end{aligned} \tag{3}$$

In this setting, *Constraints (Short)* represents our exogenous measure of firms experiencing intensified financial constraints during the financial crisis. It is calculated from the dollar amount of long-term debt due in one year in 2007, adjusted by total sales. We express this value as a percentage by multiplying it by 100.

We present the model estimates for the full sample, the sample of counties along state borders, and county pairs located within 75 km of each other in Table 1. Our findings indicate that financial constraints significantly alter the employment dynamics of establishments in states affected by the federal minimum wage change. More specifically, the results in the last column suggest that a one standard deviation change in the financial friction measure leads to a 1.1% decrease in employment.

Table 1: Financial constraints during the crisis

	(1) All counties	(2) Counties on borders	(3) Counties on borders (≤ 75 km)
	Employment (log)	Employment (log)	Employment (log)
Post × Treated	-0.004	-0.006	-0.006
	(0.005)	(0.007)	(0.007)
Post × Treated × Constraints (short)	-0.000	-0.005**	-0.005*
	(0.002)	(0.003)	(0.003)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	611,659	185,947	164,970
R-squared	0.960	0.962	0.962

Notes: This table shows regression results for Equation (3). We use *Employment (log)* as our outcome variable. All the regressions include establishment and year × firm fixed effects. We focus on three alternative samples: (i) all counties in the United States in Column (1), (ii) all counties on state borders in Column (2), and (iii) all counties on state borders whose centroids are less than 75 km apart in Column (3). Robust standard errors, clustered at the state level, are reported in parentheses below the coefficient estimates. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of these variables and their data sources.

To further illuminate the mechanism behind our findings, we examine whether the employment response varies with the level of cash holdings of the firms. Our conceptual framework predicts that financially constrained firms place a higher value on preserving cash and therefore face a greater marginal cost of absorbing higher wage bills through internal funds. Consequently, financially constrained firms with low cash holdings are expected to be more likely to adjust employment when minimum wages rise. To test this prediction, we use our preferred specification based on establishments along state borders and split the sample at the median cash-to-assets ratio measured as of 2006. We report the estimation results in Table 2.

The results align closely with these theoretical predictions. Even if the coefficients of interest are negative, Columns (1) and (2) show no statistically significant effects of the federal minimum wage change for establishments belonging to high-cash firms, consistent with their greater ability to absorb the shock internally. In contrast, Columns (3) and (4) indicate significant employment declines among low-cash firms, with the negative effect concentrated in establishments of financially constrained firms. These results support the view that financially constrained firms may decide to decrease their employment levels after a minimum wage change to preserve their cash reserves.

Table 2: Role of corporate cash holdings during the crisis

	(1)	(2)	(3)	(4)
Samples:	High Cash	High Cash	Low Cash	Low Cash
	Employment (log)	Employment (log)	Employment (log)	Employment (log)
Post×Treated	-0.010	-0.003	-0.018**	-0.010
	(0.007)	(0.007)	(0.008)	(0.011)
Post×Treated×Constraints(short)		-0.005		-0.006*
		(0.004)		(0.004)
Establishment FE	✓	✓	✓	✓
Year × Firm FE	✓	✓	✓	✓
Observations	99,627	96,534	99,574	84,966
R-squared	0.965	0.965	0.960	0.957

Notes: This table reports regression results for Equation (3). We split the sample into two groups based on cash-to-assets above and below the median and focus on establishments located along states' borders. We use *Employment (log)* as the outcome variable. All regressions include establishment and year × firm fixed effects. Robust standard errors, clustered at the state level, are reported in parentheses below the coefficient estimates. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for detailed variable definitions.

Short-term debt and other corporate characteristics. A potential issue with this approach is that while the debt maturity structure is predetermined, it may not be exogenous to other firm characteristics that influence firms' employment responses. Firms with a high level of short-term debt might systematically differ from other firms in ways that are correlated with employment decisions and are unrelated to their financial constraints. For instance, these firms might have debt rollovers more frequently because of poor operating performance or short-term strategic focus, which may also affect how firms respond to labor cost shocks.

To investigate this potential concern, we demonstrate that firms with higher and lower exposure to short-term debt are similar in a number of important observable characteristics before the crisis. Specifically, we split the firms in our sample based on the median value of our financial constraint measure and present separate summary statistics for the high- and low-exposure groups, along with their normalized differences (ND). In fact, Imbens and Wooldridge (2009) argues in favor of using a normalized difference in tests for differences between groups, rather than using t-statistics because the former measure is not affected by sample size. More specifically, they argue that two samples can be considered similar if the normalized difference falls within a ± 0.25 range.

Our findings reported in Table 3 indicate that the two groups of firms are comparable in terms of several major firm characteristics (total assets, return on assets, return on investment, asset tangibility, employment growth, cash holdings, cash volatility, and a distress indicator), supporting the validity of our identification strategy. Interestingly, we also show that these firms do not differ in their average historical frequency of debt rollovers, measured by the historical average short-term debt level for each firm (Historical Short-Term Debt).

To provide further evidence that the amount of short-term debt at the beginning of the financial crisis period is not correlated with other firm characteristics, we use it as an outcome variable and regress it against the major firm characteristics measured in 2006, while also controlling for sector fixed effects. We report the results in Table 4. Importantly, none of the coefficients is statistically significant, except for the historical use of short-term corporate debt.

Further robustness checks on the validity of our measure. To ensure that the correlation with the historical use of short-term debt does not affect our findings, we show in Column (1) of Table OA B5 that our results remain robust when

Table 3: Balance check across financially constrained groups

	High ST Debt		Low ST Debt		
	(1)	(2)	(3)	(4)	(5)
	Mean	SD	Mean	SD	ND
Log(Firm Size)	5.85	2.22	5.63	1.97	0.07
ROA	-0.10	0.54	-0.06	0.52	-0.06
ROI	-0.00	0.77	0.00	0.64	-0.01
Tangibility	0.29	0.27	0.24	0.27	0.14
Historical short-term debt	14.67	61.57	5.29	26.95	0.14
Employment growth	0.68	2.01	0.48	1.40	0.08
Cash flow volatility	0.89	2.14	1.17	2.72	-0.08
Distress	0.24	0.43	0.28	0.45	-0.06
Cash holdings	0.14	0.24	0.19	0.27	-0.12

Notes: This table presents the summary statistics of firm characteristics for the two samples created by splitting the firm sample based on the median value of firm exposure to the short-term (ST) debt maturity metrics. We report separate summary statistics for the high and low exposure groups, along with their normalized differences. See Table OA B2 for a detailed description of the variables and their sources.

we include this measure as a control and, in particular, its full interaction with both the post-period and minimum wage treatment indicators. In Column (2), we further demonstrate that the results continue to hold when we include and fully interact all other corporate characteristics presented in Table 3 with the treatment indicators. Overall, these results suggest that the other confounding corporate characteristics that we consider cannot explain our findings.

We also consider a limitation of the NETS database, that is the high imputation rate for employment information, particularly severe in small establishments. More specifically, employment information of small establishments is more likely to be imputed. For larger establishments, the imputation rate is much lower. To test the reliability of our results, we conduct a robustness check by removing establishments with three or fewer employees. Additionally, we perform another test by excluding establishments with round numbers of employees (5, 10, 100, 200, . . . , 1000), as these are more likely to represent imputed data. The results remain consistent and are reported in Columns (3) and (4).

To further test the validity of our approach, we decompose the Post variable into

Table 4: Determinants of short-term debt

	(1)	(2)
	Constraints (Short)	Constraints (Short)
Log(Firm Size)	0.190	0.061
	(0.160)	(0.178)
ROA	-0.275	-0.047
	(1.981)	(1.969)
ROI	-0.923	-0.538
	(1.157)	(1.116)
Tangibility	-1.563	-2.119
	(1.239)	(1.459)
Historical short-term debt	0.190***	0.185***
	(0.032)	(0.032)
Employment growth	-0.006	0.055
	(0.230)	(0.239)
Cash holdings	-2.105	-2.840
	(2.905)	(3.106)
Cash flow volatility	0.233	0.312
	(0.208)	(0.187)
Distress	2.352	1.514
	(1.757)	(1.697)
Constant	0.968	
	(1.037)	
Sector FE		✓
Observations	1,449	1,442
R-squared	0.350	0.384

Notes: This table presents the results of a regression analyzing corporate financial constraints during the crisis based on alternative corporate characteristics. Corporate financial constraints are measured as the share of long-term debt outstanding in 2007 that matures in 2008, adjusted by total sales. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of the variables and their sources.

individual years using a generalized difference-in-differences setting to analyze the employment dynamics around the Great Recession crisis period for financially constrained firms. More specifically, we report each yearly interaction coefficient in Column (5) of Table OA B5.

Our results show that, at the onset of the financial crisis, employment decreased substantially among establishments of financially constrained firms exposed to the federal minimum wage hike, and this effect persisted. More specifically, the triple interaction coefficient is negative and statistically significant in 2008. In addition, the triple interaction coefficients are also negative and statistically significant in 2009, 2010, and

2011. On the other hand, these dynamics support the parallel trends assumption, as none of the coefficients is statistically significant before 2008, further supporting the validity of our approach.

Finally, as a falsification test, we investigate the impact of long-term debt maturing after the end of the financial crisis, denoted as *Constraints (long)*. This variable is calculated as the sum of long-term debt due in two, three, four, and five years relative to 2007, adjusted by total sales, and expressed as a percentage by multiplying it by 100. Consistent with our prediction, the results reported in Column (6) of Table OA B5 show no evidence that *Constraints (long)* affects establishment employment levels. In fact, the interaction coefficients are approximately zero and statistically insignificant.

5.3 Firm level analysis and the financial crisis experiment

It might be supposed that firms would find it optimal to shift employment from establishments exposed to rising minimum wage levels to non-exposed establishments. This adjustment could also imply that a firm's aggregate employment changes little, if at all. However, the optimal employment level of each establishment is likely to be independent of the employment levels of other establishments, as each has its unique characteristics and optimal equilibrium. In our empirical analysis, we account for time-invariant establishment differences with establishment fixed effects. In addition, transferring employees between establishments faces several frictions that could more than offset any wage saving. These frictions include training costs, skill mismatches, employee dissatisfaction and resignations, and geographic labor constraints.

We complement our establishment-level results with a firm-level analysis to empirically investigate whether the redistribution of employees across establishments of the same firm plays a role, demonstrate the robustness of our findings, and provide deeper insight into how minimum wage policies and financial constraints influence corporate employment decisions. Specifically, we estimate the following Equation:

$$Log(Emp.)_{f,t} = \beta_0 Post_t \times Treated_f + \beta_1 Post_t \times Treated_f \times Constraints_f + \beta_3 Post_t \times Constraints_f + \eta_{s,t} + \theta_f + \epsilon_{i,t} \quad (4)$$

In this setting, *Log(Emp.)* is the natural logarithm of employees at firm f at time t, and *Treated* is measured as the share of employees of a firm f in states affected by the federal minimum wage increase. Moreover, $\eta_{s,t}$ and θ_f are respectively sector-year and firm fixed effects.

We present consistent results in Table 5. As reported in the first two columns, we continue to find a negative effect on employment for constrained firms exposed to the federal minimum wage increase. In fact, the triple interaction coefficient is negative and statistically significant. In terms of economic significance, as reported in the second column, this translates into a 2.8% decline in firm-level employment for a one standard deviation increase in both exposure to treated states and financial constraints. In Column (3), we show that our coefficients of interest are still negative and statistically significant when we control for and fully interact several corporate characteristics with the corporate level exposure to the federal minimum wage hike.[10] Finally, in Column (4), we present the dynamic effects of this set of results. In line with our establishment-level evidence, we find that the post-period interaction coefficients are negative and statistically significant, starting in the year 2008. Furthermore, these dynamics support the parallel trends assumption, as none of the coefficients is statistically significant before the shock, further supporting the validity of our approach.

5.4 Aggregate implications

With additional assumptions, we can use our establishment-level estimation results from Column (2) of Table 1 to assess the impact of corporate financial constraints

[10] We again consider total assets, return on assets, return on investment, asset tangibility, employment growth, cash holdings, the historical use of short-term debt, and the distress indicator.

Table 5: Firm employment during the crisis and financial constraints

	(1)	(2)	(3)	(4)
	Employment (log)	Employment (log)	Employment (log)	Employment (log)
Post × Treated	0.002	-0.004	-0.041	-0.038
	(0.025)	(0.023)	(0.058)	(0.058)
Post × Constraints (short)	0.000	0.000	0.001	0.001
	(0.001)	(0.001)	(0.001)	(0.001)
Post × Treated × Constraints (short)	-0.007*	-0.008**	-0.009**	
	(0.004)	(0.004)	(0.004)	
Year 2004 × Treated × Constraints (short)				0.001
				(0.002)
Year 2005 × Treated × Constraints (short)				0.003
				(0.004)
Year 2006 × Treated × Constraints (short)				-0.000
				(0.004)
Year 2007 × Treated × Constraints (short)				-0.003
				(0.004)
Year 2008 × Treated × Constraints (short)				-0.011**
				(0.005)
Year 2009 × Treated × Constraints (short)				-0.012**
				(0.005)
Year 2010 × Treated × Constraints (short)				-0.011**
				(0.005)
Year 2011 × Treated × Constraints (short)				-0.011*
				(0.006)
Firm FE	✓	✓	✓	✓
Year FE	✓	Subsumed	Subsumed	Subsumed
Year × Industry FE		✓	✓	✓
Fully interacted controls			✓	✓
Observations	12,981	12,931	10,929	10,929
R-squared	0.977	0.979	0.982	0.982

Notes: This table shows regression results for Equation (4). We use *Employment (log)* as our outcome variable. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. $^{*}p<0.10$, $^{**}p<0.05$, $^{***}p<0.01$. See Table OA B2 for a detailed description of these variables and their data sources.

economy-wide. To this end, we follow Chodorow-Reich (2014) and perform the following counterfactual exercise.

We first estimate the total impact of financial constraints on employment in our sample. In the counterfactual analysis, we assume that every establishment exposed to the federal minimum wage increase faces no financial constraints, as described by the following equation:

$$Log(Emp_{i,t})^{CF} = E[Log(Emp_{i,t})|Constraints(Short)_i = 0] = \widehat{Log(Emp_{i,t})} + \hat{\beta}_1 \times \Delta FC_i \tag{5}$$

$Log(Emp_{i,t})^{CF}$ represents the counterfactual (log) employment for establishment i at time t. $E[Log(Emp_{i,t})|Constraints(Short)_i = 0]$ is the expected (log) employment assuming the absence of a financial constraint. $\widehat{Log(Emp_{i,t})}$ is the fitted value of the employment regression. The adjustment term $\hat{\beta}_1 \times \Delta CF_i$ includes $\hat{\beta}_1$, the estimated interaction coefficient between the event indicator, the federal minimum wage treatment indicator, and the financial constraint variable, and ΔCF_i, which represents the difference in the firm's counterfactual and actual financial constraint levels. These counterfactual predictions are then averaged over establishments in states affected by the federal minimum wage change.

Figure 3 displays the results of this counterfactual analysis. The blue line shows the predicted average employment level (log) assuming no financial constraints, while the red line depicts the predicted average employment level (log) based on the actual values observed in the database.

To estimate total employment gains for our sample assuming firms face no financial constraints, we first define Employment Gains for each establishment i at time t as the difference in the counterfactual and actual employment levels:

$$\text{Employment Gains}_{i,t} = \exp(\log(Emp_{i,t})^{CF}) - \exp(\widehat{\log(Emp_{i,t})}) \tag{6}$$

We then calculate the percentage of total employment gains using the following Equation:

$$\text{Total \% Employment Gains} = \frac{\sum_{t=2007}^{2011} \sum_i \text{Employment Gains}_{i,t}}{\text{Total Employment}_{2007}} \times 100 \tag{7}$$

Total Employment$_{2007}$ represents the actual total employment level in our sample in the base year 2007. This measure allows us to calculate the percentage increase in total employment over the 2007-2011 period relative to the initial employment level in 2007, assuming no financial constraints. This exercise indicates a 1.4% increase in total employment relative to the initial aggregate employment level in our sample, abstracting from general equilibrium effects and assuming that the total employment effects represent the sum of the direct effects for each establishment.

To align our estimate with the broader establishment population, we adjust our predicted values using the weights derived from the employment distribution of the NETS data set in the establishment size categories.[11] We apply these weights when aggregating the predicted employment levels and calculating the total employment effects, scaling our sample-based estimates to better represent the overall economy, which includes public and private establishments of various sizes. This approach results in a more pronounced employment gain of 2.3%. Given that the employment gains in our empirical analysis are observed exclusively in states affected by the federal minimum wage change, these estimates have significant economic implications. We conclude that minimum wage policies can significantly impact *aggregate* employment in the presence of corporate financial constraints.

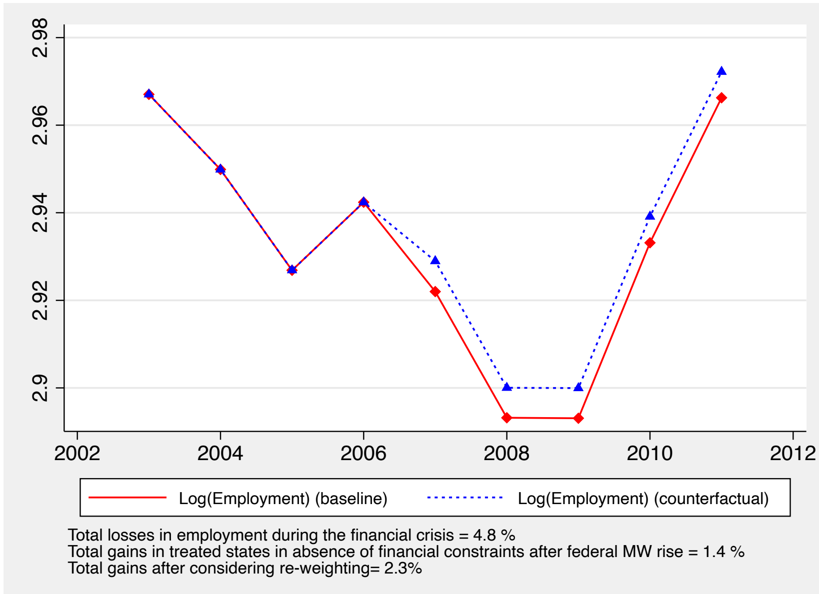
The external validity of our estimated coefficient $\hat{\beta}_1$ limits our prediction of employment gains in the overall population. Applying this estimate to all U.S. establishments is likely to be downward-biased since our sample is derived from only publicly owned establishments. However, private companies are likely to face more serious financial constraints when they experience comparable debt levels maturing during the crisis period. This suggests that our analysis provides a conservative lower-bound estimate on the total aggregate employment gains expected in the absence of financial constraints after a minimum wage rise.

6 The staggered introduction of minimum wage policies

The previous investigation focuses on a historical period marked by heightened financial constraints. As a complementary laboratory, we examine a more typical and generalizable environment in which firms are less financially constrained, allowing us

[11] Weights are reported in Table OA B6.

Figure 3: Counterfactual exercise



Notes: Figure 3 illustrates our counterfactual exercise results. The blue line represents the average Log(Employment) level assuming no financial constraints in states affected by the federal minimum wage change. The red line depicts the predicted value based on observed values. This graph focuses on establishments in states affected by the federal minimum wage change.

to better capture the behavior of unconstrained firms. More specifically, we exploit staggered changes in minimum wage states across states from 1990 to 2020. Figure OA B4 illustrates these dynamics for each state. Our empirical analysis takes advantage of this significant variability in minimum wage changes across states and periods.

Empirical analysis. In this setting, we estimate Equation (8) shown below:

$$\begin{aligned} Log(Employment)_{i,t} = {} & \beta_0 MW_{s,t} + \beta_1 MW_{s,t} \times Financial\ Constraints_{f,t-1} \\ & + \delta_{f,t} + \eta_i + \epsilon_{i,t} \end{aligned} \tag{8}$$

In this specification, MW is the effective minimum wage in state s at time t. *Financial Constraints* are measured by three alternative metrics, which are denoted by firm f in year t-1, which we link to each firm's establishments in the county. As explained in the data section, we use firm size, the WW index, and the SA index as a firm's financial constraint metrics. In this setting, the main coefficient of interest is β_1, which captures the average effect of firm financial constraints on the relationship between minimum wages and employment. On the other hand, β_0 reports the effect of the minimum wage on employment when the financial constraint variable is hypothetically equal to zero.

First, we estimate the impact of minimum wage increases on employment, excluding the financial constraint variable from the estimation process, and report the results in Table OA B7. We do not find evidence that minimum wage policies affect average establishment-level employment. None of the three coefficients of interest are statistically significant. These findings are consistent with several previous studies that analyze the effects of minimum wage changes during normal economic periods (e.g., Cengiz et al., 2019; Dustmann et al., 2022).[12]

Next, we report model estimates for the three alternative financial constraint metrics in Table 6. These results suggest that a firm's financial frictions play a crucial role in explaining employment dynamics following changes in a state's minimum wage. In fact, the coefficient of interest β_1 is positive and statistically significant when a firm's financial constraint is firm size measured by total assets, as smaller firms are more financially constrained. Alternatively, when we use the WW and SA indices to capture a firm's financial constraint, the interaction coefficients are negative and statistically significant, where higher values of these two indices represent firms with tighter financial constraints.

To assess the economic magnitude of the effects reported in Column (3), we use the standard deviation of each financial constraint measure. Starting with firm size,

[12] Figure OA B5 shows the dynamic effects of minimum wage changes on employment for establishments located along state borders in our sample. We use a two-way fixed effects (TWFE) estimator, as well as the estimators recommended by De Chaisemartin and d'Haultfoeuille (2020) and Sun and Abraham (2021) to address potential endogeneity concerns that arise when treatments are non-homogeneous and staggered. The absence of significant coefficients before the event suggests that firms in contiguous counties exhibit parallel employment trends before minimum wage changes, an essential assumption for the validity of our difference-in-differences estimator. Furthermore, we still find no evidence that changes in minimum wage policies affect establishment employment in the years that follow a minimum wage change.

the interaction coefficient implies that a one standard deviation increase in firm size attenuates the negative employment effect of a $1 increase in the minimum wage by approximately 0.85%. In contrast, the WW Index has an interaction coefficient of –0.073, implying that a one standard deviation increase in this financial constraint leads to a 0.82% decline in employment. Finally, for the SA Index, the interaction coefficient is -0.006, yielding an estimated 0.87% reduction in employment for a one standard deviation increase in this constraint measure. In summary, these estimates indicate economically meaningful effects: a one standard deviation change in the financial constraint measures reduces employment by approximately 0.8 – 0.9% per $1 increase in the minimum wage.

Motivated again by our conceptual framework, we test whether constrained firms with lower cash holdings exhibit a stronger employment response to minimum wage hikes. We split the firm sample into high and low cash-to-assets groups and re-estimate the impact of financial constraints on employment separately for each subsample using our preferred specification of establishments located along state borders.

As shown in Table 7, financial constraints significantly reduce employment only for establishments belonging to firms with limited cash holdings, while we find consistent but not statistically significant coefficients for the other sub-sample. This finding supports the notion that firms' internal liquidity buffers play a key role in shaping corporate labor adjustment decisions when facing cost shocks and external financing frictions, and further highlights the importance of precautionary savings motives for financially constrained firms in this setting.

Revenues as corporate financing mechanism. A related question in recent work is whether firms can finance the higher labor costs induced by minimum wage increases through higher prices and revenues. Consistent with this hypothesis, Rao and Risch (2025) show that small independent businesses are often able to fully offset these costs through increased revenues, likely reflecting strong local demand and some degree of price-setting power. In contrast, our evidence in Table OA B8 points to a

Table 6: Role of corporate financial constraints

	(1) All counties	(2) Counties on borders	(3) Counties on borders ($\leq$ 75 km)
	Employment (log)	Employment (log)	Employment (log)
	Panel A: Interaction with corporate size		
MW	-0.017***	-0.025***	-0.034***
	(0.006)	(0.009)	(0.011)
MW × Log(Firm Size)	0.002**	0.003***	0.004**
	(0.001)	(0.001)	(0.001)
Observations	2,283,862	690,721	610,632
R-squared	0.936	0.939	0.940
	Panel B: Interaction with the WW index		
MW	-0.020***	-0.023***	-0.033***
	(0.007)	(0.008)	(0.010)
MW × WW Index	-0.042**	-0.058***	-0.073***
	(0.018)	(0.020)	(0.025)
Observations	2,186,212	661,777	584,829
R-squared	0.938	0.940	0.941
	Panel C: Interaction with the SA index		
MW	-0.006***	-0.007**	-0.012***
	(0.002)	(0.003)	(0.004)
MW × SA Index	-0.002*	-0.005***	-0.006**
	(0.001)	(0.002)	(0.003)
Observations	2,283,862	690,721	610,632
R-squared	0.936	0.939	0.940

Notes: This table shows regression results for Equation (8). We use *Employment (log)* as our outcome variable. All the regressions include the establishment and year × firm fixed effects. We focus on three alternative samples: (i) all counties in the United States in Column (1), (ii) all counties on state borders in Column (2), and (iii) all counties on state borders whose centroids are less than 75 km apart in Column (3). Robust standard errors, clustered at the state level, are reported in parentheses below the coefficient estimates. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of our variables and their data sources.

different conclusion for publicly listed firms during our analysis period.

We estimate our baseline specification using the natural logarithm of total sales of the establishment as the outcome variable, and we find no average increase in revenues following minimum wage hikes. When we account for firm financial constraints, the results indicate that constrained firms experience a decline in sales, consistent with these firms cutting employment and output rather than raising prices. While we can

Table 7: Role of corporate cash holdings in normal economic times

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
	Panel A: Below-median cash holdings		
MW	-0.041**	-0.033**	-0.011*
	(0.016)	(0.013)	(0.006)
MW × Log(Firm Size)	0.005**		
	(0.002)		
MW × WW Index		-0.089**	
		(0.034)	
MW × SA Index			-0.008**
			(0.004)
Observations	332,473	318,970	332,473
R-squared	0.947	0.948	0.947
	Panel B: Above-median cash holdings		
MW	-0.011	-0.012	-0.004
	(0.009)	(0.009)	(0.004)
MW × Log(Firm Size)	0.001		
	(0.001)		
MW × WW Index		-0.027	
		(0.021)	
MW × SA Index			-0.002
			(0.002)
Observations	326,365	312,146	326,365
R-squared	0.941	0.942	0.941

Notes: This table shows regression results for Equation (8) for the sample of counties on borders. The dependent variable is *Employment (log)*. All regressions include establishment and year × firm fixed effects. Panel A restricts the sample to establishments belonging to firms with cash holdings below the median, while Panel B focuses on those above the median. Robust standard errors, clustered at the state level, are reported in parentheses below coefficient estimates. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for variable definitions and data sources.

not distinguish between quantities and prices, this pattern echoes our broader findings that financially constrained firms face limited scope to pass through cost shocks and instead rely more heavily on internal cash preservation and external financing. In contrast, as we report in the last Column and in line with Rao and Risch (2025), we document revenue gains for firms with greater pricing power, as measured by the Lerner Index, suggesting that price pass-through is feasible only in specific environments and under normal economic conditions.[13]

Further assessment of the validity of our analysis. One potential concern with our analysis is that financially constrained firms could exhibit a more pronounced response to minimum wage changes, not solely due to their differing financial constraints, but also due to other characteristics of firms that correlate with their financial constraints. To mitigate this concern, we incorporate firm-year fixed effects into our main specification. This revised specification enables us to account for any time-varying firm attributes and compare establishments belonging to the same firm, but having varying levels of exposure to state minimum wage policies.

In this section, we demonstrate that our findings remain robust even when we fully interact changes in minimum wages with other time-varying firm attributes. Specifically, we present evidence that our primary results hold when incorporating alternative measures of firm performance, efficiency, asset tangibility, the cash-to-asset ratio, an indicator of financial distress, and the corporate growth rate of employment over the past four years. In addition, we control for the employment beta, defined as a statistic that quantifies the relationship between the employment variability within a four-digit SIC code and the variability of total employment. This metric captures the industry's sensitivity to overall employment shocks, based on the firm's operating sector. We report these results in Table OA B9. Again, the findings are consistent with our baseline results.

Another potential concern with our analysis is that local economic conditions may be positively correlated with changes in minimum wage laws, potentially biasing our

[13] Note that our analysis differs from Rao and Risch (2025) not only because we focus on establishments of public firms, but also because we cover a longer historical period that includes several economic downturns, including the financial crisis. In addition, the results of Rao and Risch (2025) indicate that in industries such as restaurants, the least productive small firms exit after the cost shock, while surviving (more productive) firms capture this displaced demand, further boosting their revenues and profits. This result can also help explain the differences between our findings and theirs. In our setting, minimum wage increases do not, on average, trigger establishment exits among public corporations, which limits the scope of such reallocation effects.

estimates. To address this issue, our preferred specification follows a common approach in the literature by comparing counties located on opposite sides of state borders. To test this hypothesis in more detail, we estimate a regression model at the county and year level of minimum wage rates on population, employment levels, income per capita, and earnings per capita. We report our estimation results in Table OA B10 and find that minimum wage policies do not affect population size or the employment ratio. However, consistent with our prediction, we observe an increase in earnings and income per capita. To ensure against confounding factors affecting our estimates, we control for these additional variables in robustness analysis. We report the results in Table OA B11. Importantly, we show that our coefficients of interest remain consistent in sign, are statistically significant, and are within one standard deviation of our baseline specification.

To further assess the robustness of our findings, we examine whether our baseline results hold after controlling for other significant state policies. To do so, we gather data on the unemployment insurance (UI) benefit schedules for each state from the U.S. Department of Labor's Significant Provisions of State UI Laws. We then follow Guo et al. (2024) and calculate the overall UI benefit level in a given state and year by multiplying the maximum weekly benefit amount by the maximum duration of benefits provided under each state's regular UI program.[14] Additionally, we collect information on the rates of union membership (UMR) by state from the Current Population Survey (CPS)[15] and state corporate tax changes, median state income tax rate, and top income tax rate from Baker et al. (2025). We include these added state-level controls in Equation (8). The outcomes are reported in Table OA B12, and this evidence shows that our primary findings are robust to controlling for these alternative policies adopted at the state level.

In an additional test, we follow Dube et al. (2010) and present an alternative iden-

[14] To validate this measure, Agrawal and Matsa (2013) and Hsu et al. (2018) show that changes in this variable are strongly correlated with higher aggregate state UI expenditures.

[15] For details, see `https://www.unionstats.com/` (this reference was last accessed on December 2, 2025).

tification strategy. More specifically, we opt for a method by which each county in the border region is carefully matched with all the other possible county pairs in other states. This strategy requires a change in the analyzed samples, as the establishments can be represented multiple times for each county pair they belong to. Furthermore, to ensure a meaningful comparison of contiguous counties in terms of their employment trends, we incorporate control variables for pair-by-year fixed effects. By introducing these fixed effects, we expect to capture the shared economic trends within the local labor market. The results of estimating Equation (8) are presented in Table OA B13. Notably, all the findings in this table exhibit signs and significance levels consistent with our baseline results. In addition, we undertake a more advanced approach recently proposed by Jha et al. (2025). They demonstrate that a positive bias arises in a county-pair specification when pairs are formed using contiguous counties across state lines, but where the pairs are in different commuting zones. Reassuringly, our results reported in Table OA B14 remain robust when using multi-state commuting zones to create bordering county pairs, which they argue offers a superior definition of local economic areas.

Again, we consider the limitation of the NETS database is its high imputation rate for employment information, particularly severe in small establishments. To test the reliability of our results, we again conduct a robustness check by removing establishments with fewer than three employees. In addition, we perform another test by excluding establishments with round numbers of employees (5, 10, 100, 200, . . . , 1000), as these are also more likely to represent imputed data. The results remain consistent and are reported in Tables OA B15 and OA B16.

Alternative measures of financial constraints. While our measures of corporate financial constraints are widely used in the corporate finance literature, we acknowledge that no single proxy can perfectly capture a firm's underlying degree of financial constraint. Therefore, we investigate whether our results hold when considering alternative measures of corporate financial constraints.

We start by considering the first financial constraint index commonly used in the literature, namely the Kaplan-Zingales (KZ) index (Kaplan and Zingales, 1997). The KZ index is found to be uncorrelated with the other commonly used financial constraint indexes (Farre-Mensa and Ljungqvist, 2016), and to underperform them (Whited and Wu, 2006; Hadlock and Pierce, 2010). We estimate Equation (8) using the KZ measure and report the results in Column (1) of Table OA B17. Not surprisingly, the triple interaction coefficients are close to zero and are not statistically significant in various alternative specifications.

To further assess the robustness of our findings, we build a composite financial constraint index by adopting a methodology similar to that presented by Bartram et al. (2022). To do so, we consider a method that relies on the rankings of firms based on the KZ, SA, and WW indices, as well as on firm size. Specifically, we consider whether the firm has above-median annual values for the first three indices and below-median annual values for firm size. If the majority of these indicators suggest that a firm is financially constrained, then we assign a value of one to the composite indicator; otherwise, we assign a zero value.

We present the results in Column (2) of Table OA B17; according to the estimate, firms categorized as financially constrained on average decrease their employment by 1 percentage point after a one-dollar increase in the minimum wage. Furthermore, our results continue to hold also when we consider a composite index based on the first principal component of these four alternative indexes of corporate financial frictions in Column (3) of Table OA B17. In terms of magnitude, one standard deviation increase in this combined financial constraint index decreases employment by 0.5 percentage points after a one-dollar increase in the minimum wage.

We next demonstrate the robustness of our results when considering an alternative financial constraint measure based on textual analysis of firms' annual 10-K filings. To do so, we use the dictionary of 'constraining' words proposed in Bodnaruk et al. (2015). They show that the frequency of these words exhibits a very low correlation with traditional measures of financial constraints, yet it predicts subsequent liquidity events better than widely used financial constraint indexes.

We take the natural logarithm of the word frequency of this dictionary and then control for differences in the documents' length.[16] Subsequently, we estimate Equation (8) and present the results in Column (4) of Table OA B17. In line with our baseline results, we find that after a minimum wage rise, establishments belonging to financially constrained firms reduce their workforce. More specifically, based on the estimated coefficient, after a one-dollar rise in the minimum wage rate, a 1% change in the financial constraint index based on textual analysis leads to a 1.2% decrease in employment.

Finally, we propose two additional financial constraint measures: indicators for whether a firm issues bonds or obtains syndicated loans. These measures serve as clear, objective indicators of a firm's access to external capital. Issuing bonds or obtaining syndicated loans generally requires a certain minimum level of creditworthiness and market confidence (Faulkender and Petersen, 2006; Campello et al., 2010; Acharya et al., 2013). Firms that secure these financing sources are typically less financially constrained because they can tap into external funds at favorable rates. In contrast, firms that do not issue bonds or obtain syndicated loans are likely to face tighter credit conditions and higher financing costs.

Our results hold for both measures. More specifically, in terms of economic magnitude, the interaction coefficient in Column (5) of Table OA B17 indicates that establishments owned by firms that did not obtain a syndicated loan reduce employment by approximately 0.3% for every $1 minimum wage increase, compared to those that did. Similarly, the coefficient in Column (6) implies that firms without bond issues reduce employment by 0.7% for a $1 increase in the minimum wage. These effects are in line with the magnitudes reported using the earlier financial constraint measures such as the WW and SA indexes. These overall results reinforce the conclusion that limited access to external financing worsens the negative employment effects of higher labor

[16] This information is available starting from the year 1993 and has been made available by W. McDonald on his website: https://sraf.nd.edu/sec-edgar-data/lm$_$10x$_$summaries/.

costs.

6.1 Employment responses of financially unconstrained firms

We use our three main firm financial constraint measures to analyze the employment effects of minimum wage changes across terciles of these metrics to more fully understand the employment responses of financially constrained and non-financially constrained firms. More specifically, we modify Equation (8) and interact the minimum wage rate with indicators for each financial constraint tercile, using the first tercile as the reference group. In this context, the minimum wage coefficient represents the effect of the minimum wage rate when the financial constraint variables are hypothetically equal to zero.

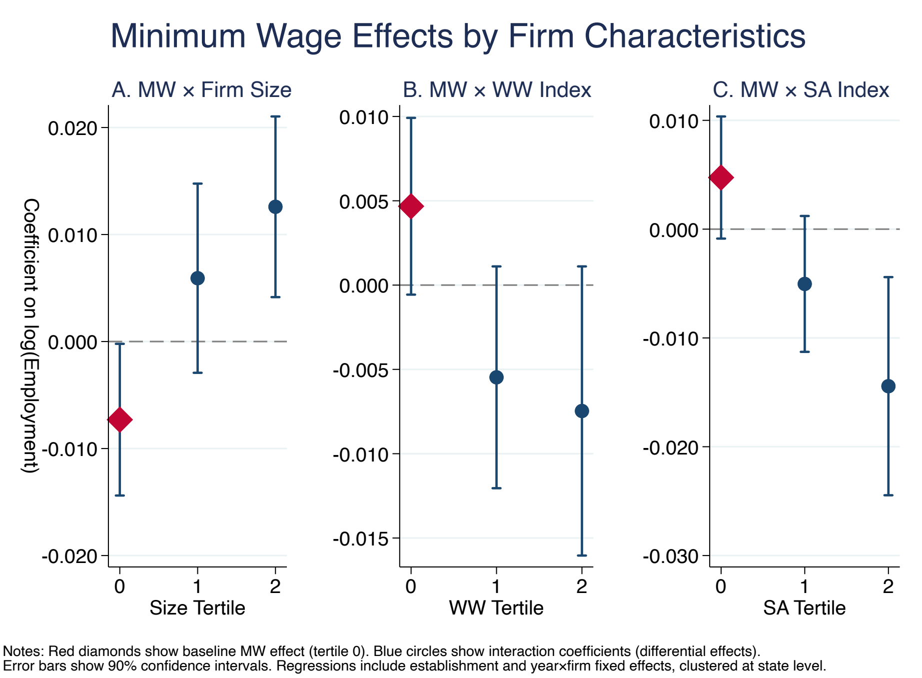
We plot the coefficients of interest and the relative confidence intervals in Figure 4. We find that the effects are roughly linear across the financial constraint range and are substantial for the most constrained firms, which exhibit adverse and economically meaningful employment effects, with two of the three coefficient estimates being statistically significant at the 5% level. On the other hand, in line with our conceptual framework, these findings also suggest that unconstrained firms are better positioned to absorb higher labor costs, benefit from the positive aspects of minimum wage increases, and expand employment.

To provide more direct evidence on how establishments owned by unconstrained firms increase their employment levels following an increase in the minimum wage rate, we estimate the following Equation:

$$Log(Employment)_{i,t} = \beta_0 MW_{s,t} + \beta_1 MW_{s,t} \times Unconstrained_{f,t-1} + \delta_{f,t} + \eta_i + \epsilon_{i,t} \tag{9}$$

In this setting, *Unconstrained* is an indicator variable that equals one for firms in the least financially constrained tertile of the firms' financial constraint distribution, and

Figure 4: Marginal effects of minimum wage changes across the financial constraints distribution



Notes: Figure 4 plots the estimated coefficients of the interaction terms between the minimum wage rate and indicators for each tercile of the financial-constraint measures, using the first tercile as the reference group. The coefficient on the minimum wage variable represents the effect when the financial-constraint measure is hypothetically equal to zero (denoted as "0" in the figure). We focus on establishments located along states' borders. Vertical lines represent 90% confidence intervals.

zero otherwise. β_1 is the coefficient of interest that represents the impact of minimum wage policies on employment in establishments owned by the least constrained firms. This is based on our preferred set of fixed effects and the sample of establishments in counties located along state borders.

We report the results in Table 8. We show that the interaction coefficients for the three alternative financial constraint measures are positive, and two of them are statistically significant. More specifically, the interaction term for unconstrained firms ranges from about 0.006 to 0.008, implying that the effect of a one-unit increase in the minimum wage is approximately a 0.6-0.8 percentage point rise in employment. In other words, if the minimum wage increases by one dollar, unconstrained establishments exhibit roughly a 0.6–0.8% increase in employment compared to more constrained firms.

Table 8: Establishments of unconstrained firms and employment

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
MW	-0.003	-0.001	-0.002
	(0.003)	(0.003)	(0.003)
MW × Unconstrained - Firm Size	0.008*		
	(0.004)		
MW × Unconstrained - WW Index		0.006	
		(0.004)	
MW × Unconstrained - SA Index			0.007*
			(0.004)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	690,721	661,777	690,721
R-squared	0.939	0.940	0.939

Notes: This table shows regression results for Equation (9). We use *Employment (log)* as our outcome variable. All the regressions include the establishment and year × firm fixed effects. We use the sample of establishments in counties located along state borders. Robust standard errors, clustered at the state level, are reported in parentheses below the coefficient estimates. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of our variables and their data sources.

How a minimum wage change can lead to higher employment? We use alternative metrics for local labor market conditions to more precisely test whether and when the establishments of unconstrained firms increase employment following a minimum wage rise. The goal of this exercise is to assess and evaluate several competing economic arguments proposed in the minimum wage literature to explain the observed increase in employment after a minimum wage hike. More specifically, we investigate and empirically test the predictions of three alternative theoretical frameworks found in the literature that could potentially explain this employment rise: (1) a job search model, (2) a labor market frictions model, and (3) employer monopsony.

Building on classic job search theory, we consider that an increase in the minimum wage can attract previously unemployed workers, thereby expanding the local labor supply. A job search model applied to these individuals could lead to higher employment levels because it can motivate firms to do more advertising of job vacancies, given the potential increase in supply and, thus, be able to fill more jobs (Burdett and

Mortensen, 1998; Flinn, 2006; Giuliano, 2013; Manning, 2021). A larger pool of workers provides corporations with more options to select candidates whose skills, motivation, and qualifications best meet their needs. This, in turn, reduces the time and costs associated with recruiting and training new workers (owing to better matches), while increasing the relative marginal benefits of these workers and enabling companies to scale their workforce more quickly in response to changes in demand.

To test this hypothesis, we identify labor markets with a potentially large supply of minimum wage workers. Using data from the 1990 Census, we gathered information on individuals of working age who have low education levels (no college education) and are unemployed or are not actively seeking employment at the beginning of our sample period. We then aggregate this information to the state level as county-level data is unavailable.

Once we obtain this variable, we divide it into two equal groups, representing high and low potential labor supply, and separately estimate Equation (9) for each group using our preferred sample of counties along state borders. The results are presented in Table OA B18. We find that unconstrained firms tended to increase their employment levels in areas with a larger local pool of potential minimum wage workers measured at the beginning of our sample period (above the median sample). In contrast, we do not find any significant effects in local markets with smaller pools of potential minimum wage workers (below the median sample).

Similarly, we consider that a minimum wage increase could reduce employee turnover, leading to an increase in a firm's workforce. If this is the case, then we expect stronger labor market responses from establishments that belong to unconstrained firms in localities where employee retention is more challenging or where employee turnover is higher. The evidence from these experiments aligns well with the models of labor markets facing search frictions, where a higher minimum wage reduces job-to-job transitions by reducing the arrival rate of better-paying job offers (Portugal and Cardoso, 2006; Brochu and Green, 2013; Dube et al., 2016, 2019).

To test the employee turnover hypothesis, we collect state-level data on job hiring, openings, and quits from the Bureau of Labor Statistics (BLS). Using this information, we construct two measures: a hiring gap, defined as the ratio of job openings minus job hires scaled by job openings, and a quit-to-hire ratio, defined as the ratio of quits to job hires.[17] We use these variables to split the database into two subsamples based on their median values, and separately estimate Equation (9) for the two groups using our preferred sample of counties along state borders.

As shown in Tables OA B19 and OA B20, we find that unconstrained firms increase their employment when their establishments are located in areas with higher hiring gaps and quit-to-hire ratios (above the sample medians). These results suggest that for financially unconstrained firms: (1) minimum wage hikes facilitate the matching process between firms and employees, and (2) minimum wage increases reduce employee turnover in local labor markets where employee retention is more challenging.

An alternative theory that explains an increase in employment after a minimum wage increase is that such increases can stimulate employment growth in monopsonistic labor markets, markets where employers have significant market power in setting wages due to limited competition for low-wage workers. In these markets, firms pay wages below the value of labor's marginal product. A higher minimum wage restricts firms' ability to pay below the competitive wage rate, which can lead them to expand employment after a minimum wage increase (Sullivan, 1989; Ransom, 1993). Recent evidence from Azar et al. (2024) supports a positive effect of minimum wage increases in concentrated labor markets.

To test this monopsony hypothesis, we construct a Herfindahl-Hirschman Index (HHI) at the county-industry labor market level using our establishment-level employment information and then estimate Equation (9) separately for the high- and low-HHI industry-county labor market concentration levels. The results are presented in Table OA B21. We do not observe significant positive effects for establishments located in counties with concentrated local labor market levels (above the median HHI). Interest-

[17] We use information as of 2000, which is the first year that this information is available.

ingly, these results suggest that establishments in counties with less concentrated labor markets experience larger positive effects. This result may seem counterintuitive under monopsony-based models, which predict stronger employment gains when employers' wage-setting power is constrained. However, in more competitive labor markets, firms may be better able to fill vacancies when the minimum wage rises, especially if turnover declines and matching efficiency improves. Moreover, unconstrained firms in competitive areas may strategically expand to attract better workers or capture market share from financially weaker competitors that are less able to absorb the higher labor costs.

Overall, these results confirm that financially unconstrained firms increase employment following minimum wage hikes and that this positive effect can be explained by a rise in local labor supply and reduced employee turnover.

Overall employment effects. In a general equilibrium context, employment can shift from firms facing significant financial constraints to firms with strong balance sheets. The magnitude of this labor reallocation depends on various factors, including local labor market conditions, the capacity of non-financially constrained firms to offer higher product prices, and the substitutability of goods among these establishments (Chodorow-Reich, 2014; Gilchrist et al., 2017; Mian and Sufi, 2014).

To test whether the labor reallocation that we uncover can fully explain the zero effect on employment after a minimum wage rise, we adopt the methodology outlined by Dube et al. (2010), which employs county-level regressions, as reported in Equation (10).

$$\begin{aligned} Log(County\ Emp.)_{j,t} = {} & \beta_0 MW_{s,t} + \beta_2 Constraints_{j,t-1} \\ & + \beta_1 MW_{s,t} \times Constraints_{j,t-1} + \delta_j + \eta_c + \epsilon_{i,t} \end{aligned} \tag{10}$$

More specifically, we measure county-level (j) employment as the total employment of all establishments within a county that is available in our sample. As in our earlier analysis, the county-level financial constraint measures are the employment-weighted average values derived from the three alternative measures of financial constraints across all establishments in our sample within each county.

The findings in Table OA B22 echo our baseline results, revealing that counties with greater exposure to establishments of the firms with a high level of financial constraints face greater aggregate employment declines after a rise in the minimum wage. From the results reported in the last column of this table, we find that a one standard deviation rise in one of the alternative financial constraint indexes reduces total employment by approximately 1% of average employment following a one-unit minimum wage hike. Interestingly, the coefficients that capture the effect of financial constraints on the local economy when the minimum wage is hypothetically close to zero are positive and statistically significant. This result suggests that in the absence of minimum wage pressures, constrained firms might rely more heavily on labor, which is cheaper, rather than on more costly capital investments.[18]

6.2 Firm-level analysis and minimum wage rates

In this section, we explore how corporate exposure to shifts in the minimum wage influence overall employment decisions, financial performance, and corporate policies. This investigation allows us to gain a deeper understanding of how minimum wage policies impact the employment choices of financially constrained firms.

Corporate employment. To empirically investigate whether the redistribution of employees across establishments of the same firm plays a role and the robustness of our results, we turn to the firm-level regressions. For this purpose, we measure corporate dollar exposure to changes in minimum wage policies for firm f at time t

[18] To examine aggregate employment effects, we focus on establishments of public corporations. These effects should be even stronger when including all establishments, as private firms are more likely to face financial constraints. To test this, we use the full NETS database, which includes private establishments and their parent firms. As reported in Table OA B23, establishments of private firms reduce employment by about 0.4% following a $1 increase in the minimum wage, consistent with our hypothesis.

using a weighted average approach, as reported in the following Equation:

$$MW\ Exposure_{f,t} = \sum_{n=1}^{51} Share\ Employees_{c,s,t} \times MW_{s,t} \tag{11}$$

In this setting, *Share Employees* is the share of employees of company f in state s at time t and MW is the effective minimum wage in state s at time t.

We next estimate the following Equation:

$$\begin{aligned} Log(Emp.)_{f,t} = \beta_0\ MW\ Exposure_{f,t} + \beta_1\ MW\ Exposure_{f,t} \times Financial\ Constraints_{f,t-1} \\ + \beta_3 Financial\ Constraints_{f,t-1} + \eta_{s,t} + \theta_f + \epsilon_{i,t} \end{aligned} \tag{12}$$

$\eta_{s,t}$ and θ_f are, respectively, sector-year and firm fixed effects. One limitation of this specification is that we are unable to control for firm-year fixed effects, which are co-linear with our other dependent variables of interest. Additionally, we cannot use a border discontinuity approach since our treatment variable has been constructed by considering the entire distribution of establishments across U.S. states, rather than focusing solely on corporate headquarters.

We report the model estimates in Table 9. These results are consistent with our previous findings at the establishment level. We continue to find that financially constrained firms decrease their employment after increases in the minimum wage. In terms of magnitude, this effect is even greater than the results that we documented earlier at the establishment level. More specifically, we find that a one standard deviation increase in financial constraint decreases overall corporate employment by approximately 2% after a $1 increase in the minimum wage exposure variable.[19]

Table 9: Firm level analysis - Employment and minimum wage rates

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
MW Exposure	-0.043***	-0.017	0.015
	(0.012)	(0.011)	(0.010)
Log(Firm Size)	0.251***		
	(0.019)		
MW Exposure × Log(Firm Size)	0.008***		
	(0.002)		
WW Index		-0.913***	
		(0.295)	
MW Exposure × WW Index		-0.152***	
		(0.037)	
SA Index			-0.121***
			(0.017)
MW Exposure × SA Index			-0.018***
			(0.002)
Firm FE	✓	✓	✓
Year × Industry FE	✓	✓	✓
Observations	54,829	50,158	54,829
R-squared	0.961	0.949	0.953

Notes: This table shows regression estimates for Equation (12). We use *Employment (log)* as our outcome variable. All the regressions include firm and year × industry fixed effects. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of our variables and their data sources.

Corporate policies and performance. We use this setting to investigate how exposure to minimum wage policies affects corporate financial policies and performance. According to our theoretical framework, we expect minimum wage policies to exert pressure on firms' internal resources. This pressure can force firms to reduce

[19] A potential issue with employment information reported in Compustat is its tendency to aggregate the total number of employees at the global level. To examine whether this impacts our findings, we exclude multinational firms from the sample and re-estimate Equation (12). Specifically, a firm is classified as multinational if it reports non-zero foreign income in the previous three years. Alternatively, we define a firm as multinational if at least 5% of its sales were generated from outside its home country, based on the Compustat Geographic Segment database (Nimier-David et al., 2023). We find consistent results in Tables OA B24 and OA B25.

expenditures and to seek external resources to address their short-term liquidity needs and longer-term financial requirements.

To investigate whether these dynamics hold, we create additional corporate-level variables and estimate the following Equation:

$$Outcome_{f,t} = \beta\ MW\ Exposure_{f,t} + \eta_{s,t} + \theta_f + \epsilon_{i,t} \tag{13}$$

In this setting, $\eta_{s,t}$ and θ_f are respectively sector-year and firm fixed effects. *Outcome* alternatively represents one of the following variables: Liquidity, Leverage (Short Term), Leverage (Long Term), R&D, and Capital Expenditures. These variables are defined in detail in Table OA B2.

We report the results in Table 10. In the first column, we examine how corporate exposure to minimum wage policies affects liquidity buffers. We find that a one-dollar increase in the minimum wage leads to a decrease in liquidity holdings of 0.008 units. This corresponds to a 4% decrease relative to average value of the outcome variable. Therefore, the effect is also economically significant.

One plausible explanation for the dynamics of this outcome variable is that minimum wage policies cause certain firms to experience short-term liquidity needs. Because exposure to minimum wage policies affects availability to a firm's internal resources, companies subject to these policies may seek additional external financing to address funding needs, meet other financial obligations, or preserve cash reserves. This, in turn, can account for the increase in long-term leverage reported in Column (2) (Custódio et al., 2013). Indeed, a firm's leverage coefficient is statistically significant, with a one-unit increase in the minimum wage measure associated with a 0.01 increase in its debt-to-assets ratio, which is a 4% change with respect to its average value in our sample. On the other hand, we do not find any effect on short-term debt (Column 3), as firms mainly rely on cash instead.

In Columns (4) and (5), we investigate the impact of exposure to minimum wage policies on corporate investment decisions. We find that a one-unit increase in the minimum wage exposure measure is associated with a half-basis point decrease in R&D expenditures, corresponding to a 5.5% decline in average R&D expenditures. Additionally, we also observe that exposure to minimum wage policies reduces capital investment. However, this coefficient is not statistically significant at conventional levels.

Finally, in the last column, we also consider how exposure to minimum wage policies affects firm performance, using return to assets as the outcome variable of interest. Consistent with our establishment-level analysis, we do not find any evidence that exposure to minimum wage policies affects average corporate performance.

Table 10: Corporate policies and operating performance

	(1)	(2)	(3)	(4)	(5)	(6)
	Liquidity	Leverage (Long)	Leverage (Short)	R&D	Capital Exp.	ROA
MW Exposure	-0.008**	0.010**	-0.001	-0.004**	-0.001	-0.001
	(0.004)	(0.004)	(0.002)	(0.002)	(0.001)	(0.003)
Firm FE	✓	✓	✓	✓	✓	✓
Year × Industry FE	✓	✓	✓	✓	✓	✓
Observations	48,422	48,308	48,435	28,616	48,055	48,425
R-squared	0.599	0.594	0.561	0.763	0.609	0.668

Notes: This table shows regression results for Equation (13). We use alternative outcome variables as reported in the second row. All the regressions include firm and year × industry fixed effects. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of our variables and their data sources.

The combined change in liquidity and long-term debt are roughly consistent with the implied increase in labor costs. Based on our own estimation, a $1 increase in the minimum wage raises total wages by about $929 per employee.[20] Given that the median firm in our sample employs roughly 1,300 workers, this corresponds to a median wage-bill increase of approximately $1.2 million per firm. To compare these wage increases with balance-sheet adjustments, we use our regression estimates showing

[20] Estimation results are reported in Table OA B26, as well an explanation of our approach.

that firms reduce liquidity holdings and increase long-term debt following minimum-wage exposure. Because these variables are expressed as ratios scaled by lagged total assets, we convert the coefficients into dollar terms using the median level of assets ($254.47 million). The implied adjustments are a decrease of about $1.9 million in cash and short term investment holdings and an increase of about $2.5 million in long-term debt.

Overall, these findings indicate that firms finance the wage-bill increase primarily by drawing down cash reserves and expanding long-term debt to finance the higher wage bill in later periods and for other future contingencies. Considering our inability to apply a border discontinuity approach in this setting, these estimates should not be interpreted as causal, but rather as roughly economically consistent with the magnitude of the observed wage adjustments.[21]

Liquidity, debt, and financial constraints. We explicitly test whether unconstrained firms are the ones using liquidity buffers and long-term debt to finance higher wage costs. Specifically, we use firm-level liquidity levels as the outcome variable and interact minimum wage exposure with terciles of our financial constraint measures.

The results reported in Panel A of Table OA B27 confirm the expected pattern: unconstrained firms reduce liquidity holdings following minimum wage increases more than constrained ones. More specifically, across all specifications, unconstrained firms, those in the top size tercile or with low values of the WW and SA indices, reduce their liquidity levels after minimum wage hikes, consistent with the use of internal funds to cover higher labor expenses.

We turn to long-term debt as outcome variable in Panel B of Table OA B27, and

[21] In addition, the estimate of the wage bill should also be interpreted with caution. The average increase of nearly one thousand dollars, estimated using data from the CBP, may be higher for public corporations. This is because they are expected to respond more strongly to a minimum wage increase given their greater transparency, heightened reputation concerns, and more rigid pay structures. The latter feature has been shown to generate spillovers, including across establishments within the same firm, that are roughly 30% larger, as firms seek to maintain wage differentials to preserve incentive structures across the organization (Hjort et al., 2026; Caires, 2025).

we document similar patterns. Using the WW and SA indices, we find that, following a minimum-wage increase, financially constrained firms experience a smaller increase in debt than unconstrained firms. This is reflected in the positive and statistically significant coefficient relative to the baseline category, where these variables are hypothetically equal to zero, indicating no constraints. On the other hand, we find weak evidence that unconstrained firms increase their use of long-term debt more than constrained firms in response to minimum-wage exposure when considering the size variable (as the coefficient is positive and large but not statistically significant at the conventional level). This result suggests that larger firms mainly rely on liquidity holdings to accommodate higher labor costs.

7 Conclusions

An extensive literature in economics analyzes the effect of minimum wage policies on employment and finds mixed results (e.g., Card and Krueger, 1994; Cengiz et al., 2019; Clemens and Wither, 2019). Our paper contributes to this literature by analyzing whether corporate financial constraints can help further explain the employment dynamics at heterogeneous firms following minimum wage changes.

To advance our research, we collect information on the individual establishments that belong to publicly traded U.S. firms and the corporate balance sheet characteristics of the firms that own them. We first provide causal evidence about the relationship between a rise in the minimum wage, financial frictions, and employment using a unique quasi-experimental setting. Our analysis exploits the federal minimum wage increase during the 2007-2008 financial crisis and the heterogeneity across firms in long-term debt levels maturing during the same crisis period. Using this clean identification strategy, we find that financially constrained firms in states affected by federal minimum wage regulations suffer further declines in their employment levels. In addition, we provide evidence that the precautionary savings motives of financially constrained firms helps explain these patterns.

Next, exploiting the staggered introduction of minimum wage policies across states, we do not to find any evidence that minimum wage increases, on average, affect employment. However, in line with our research hypothesis, we document that this average effect masks important corporate-level heterogeneity. When we interact a rise in minimum wages with alternative measures of corporate financial frictions, we find a large negative effect for financially constrained firms. Interestingly, we uncover positive effects at establishments of financially unconstrained firms and provide suggestive evidence that these results can be potentially explained by the fact that minimum wage policies increase job search activity among potential minimum wage workers and improve the matching between employers and employees.

Our findings provide new insight into how minimum wage policies interact with firm' financial strategies, performance, and subsequent employment decisions. They highlight the central role of financial constraints and access to external capital in shaping employment responses to minimum wage changes, helping to explain the heterogeneous effects documented in the literature. These results also offer guidance to policy makers by clarifying how firm-specific characteristics and the macroeconomic environment influence the effectiveness of minimum wage interventions, enabling more targeted and efficient policy design.

These conclusions are subject to important caveats. Our identification rests on quasi-experimental variation that, while plausibly exogenous, cannot fully eliminate concerns about unobserved heterogeneity correlated with both corporate financial constraints and exposure to minimum wage changes. In addition, the proxies for corporate financial constraints used in the corporate finance literature are imperfect along multiple dimensions. We mitigate these issues through two complementary identification strategies, alternative measures of constraints, and an extensive set of robustness checks, but the inherent limitations of the empirical exercise should be borne in mind when interpreting our findings.

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Online Appendix

This Appendix is for Online Publication (OA) and provides further details on our conceptual framework, the data, and the results of the paper "Corporate Financial Constraints, Minimum Wage Policies, and Employment". More specifically, Online Appendix A presents a simple model that formalizes our hypothesis, Online Appendix B provides additional figures and tables, and Online Appendix C describes the construction of our final sample.

Online Appendix A Conceptual framework

To better frame our research hypothesis, we develop a simple model in which firms maximize profits subject to a production function $Y_i = A_i F(L_i)$, where Y_i is the output, A_i is the firm-specific productivity, and L_i is labor input. The average employee wage is denoted by w_{ave}.[22]

For unconstrained firms, the profit maximization problem is captured by Equation (OA1):

$$\max_{L_i} \pi_i = pA_iF(L_i) - w_{ave}L_i \tag{OA1}$$

The first-order condition (FOC) is given by Equation (OA2):

$$pA_iF_L(L_i) = w_{ave} \tag{OA2}$$

A minimum wage law increases the average wage of firm i from w_{ave} to w'_{ave}. Therefore, the standard model predicts a decrease in employment. However, within this framework, it is important to note that employment could be unaffected if, for example, firm productivity or output prices increase sufficiently, that is whether $\frac{A'_i}{A_i} > \frac{w'_{ave}}{w_{ave}}$ or $\frac{p'}{p} > \frac{w'_{ave}}{w_{ave}}$. These two solutions could, for example, be attributed to efficiency wage

[22] The use of capital in the production function and any potential substitution effect of labor with capital are intentionally not considered to maintain model simplicity. However, including capital does not affect our conclusions.

effects ($Y_i = A_i(w_{ave})F(L_i)$, where $A_i(w_{ave})$ is increasing in w_{ave}) or aggregate demand effects ($p = p(w_{ave})$, where $\frac{dp}{dw_{ave}} > 0$).

We next introduce financial constraints in the model by recognizing that a firm's ability to pay employee salaries is limited by its available financial resources and profits:

$$w_{ave}L_i \leq B_i + \pi_i \tag{OA3}$$

where B_i represents a firm's internal financial resources. This constraint highlights the crucial role internal resources play in determining a firm's ability to pay employee wages.

For financially constrained firms, the optimization problem becomes:

$$\max_{L_i} \pi_i = pA_iF(L_i) - w_{ave}L_i \tag{OA4}$$

subject to $w_{ave}L_i \leq B_i + \pi_i$.

To solve this constrained optimization problem, we specify the following Lagrangian:

$$\mathcal{L} = pA_iF(L_i) - w_{ave}L_i + \lambda(B_i + pA_iF(L_i) - w_{ave}L_i - w_{ave}L_i) \tag{OA5}$$

where $\lambda \geq 0$ is the Lagrange multiplier, which in this setting represents the shadow cost of the financial constraint.

The FOC given a firm's financial constraint is reported in Equation (OA6):

$$\frac{\partial \mathcal{L}}{\partial L_i} = pA_iF_L(L_i)(1+\lambda) - w_{ave}(1+2\lambda) = 0 \tag{OA6}$$

Rearranging the FOC, we obtain the following Equation:

$$pA_iF_L(L_i) = w_{ave}\frac{1+2\lambda}{1+\lambda} \tag{OA7}$$

This condition shows that constrained firms behave as if they face a higher effective wage rate of $w_{ave}\frac{1+2\lambda}{1+\lambda}$, which exceeds w_{ave} for any positive λ. The shadow cost λ, therefore, captures the additional value of relaxing the financial constraint.

When the minimum wage increases the average wage from w_{ave} to w'_{ave}, it directly affects the firm's internal resources. For this reason, the constraint is adjusted to:

$$w'_{ave} L_i \leq B_i + \pi_i - \Delta_i \tag{OA8}$$

$\Delta_i = (w'_{ave} - w_{ave})L_i$ represents the reduction in internal resources due to the wage increase. This shows how the minimum wage increase reduces a firm's internal resources, potentially tightening its financial constraints.

For constrained firms, this leads to a FOC:

$$pA_i F_L(L_i) = w'_{ave} \frac{1 + 2\lambda'}{1 + \lambda'} \tag{OA9}$$

where λ' is generally going to be larger than the original λ due to the tighter financial constraint, further increasing the effective wage rate faced by the firm.

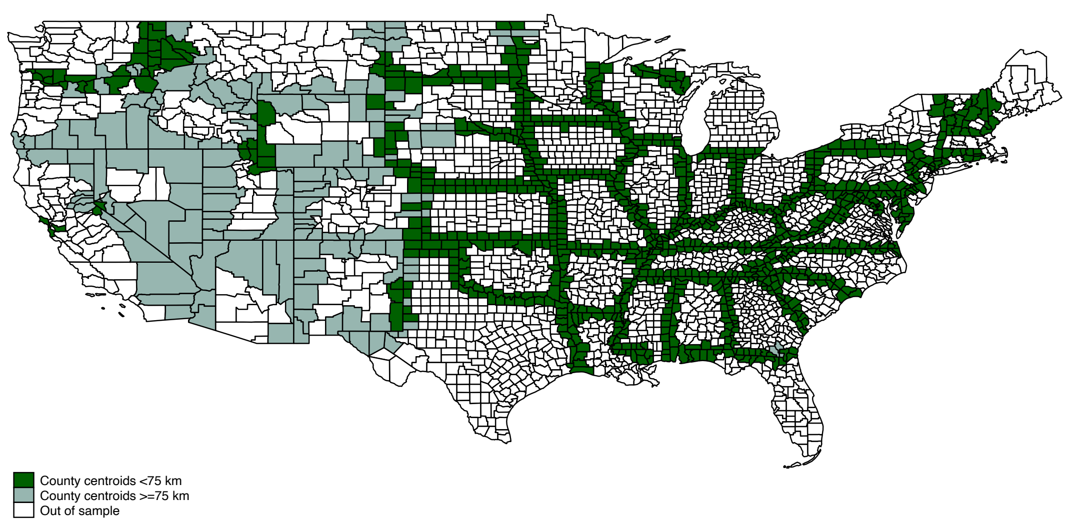
This simple model suggests that unconstrained firms can adjust their employment levels, following minimum wage increases. However, financially constrained firms face an additional effective cost due to their financial constraint, captured by the shadow cost λ. Reducing internal resources (Δ_i) further tightens the operating constraint of financially constrained firms, potentially exacerbating their reduction in employment by increasing λ.

Our empirical finding of no average effect on employment can still be reconciled through the mechanisms outlined in the original model. The key distinction is that financially constrained firms can exhibit stronger negative responses due to the added pressure exerted by their financial constraints, as captured by the shadow cost λ, leading to a potentially larger firm response following a minimum wage increase.

Online Appendix B Figures

Counties along the state border. We follow the previous literature and focus our empirical analysis on geographically adjacent counties located along states' borders to ensure that omitted local economic variables do not affect our results (Card and Krueger, 1994; Huang, 2008). To further ensure the validity of our analysis, we take into account the concern raised in Dube et al. (2010) that counties on state borders in the western U.S. are much larger and irregular in shape, and as such, they may not always share the same local economic conditions. To address this concern, Dube et al. (2010) further investigates counties on opposite sides of a state border by requiring their centroids to be within 75 km of each other. This 75 km distance cutoff is determined through a data-driven randomization inference procedure, which minimizes the mean squared error of the estimator (Dube et al., 2010). We highlight these counties by shading them in dark green in Figure OA B1.

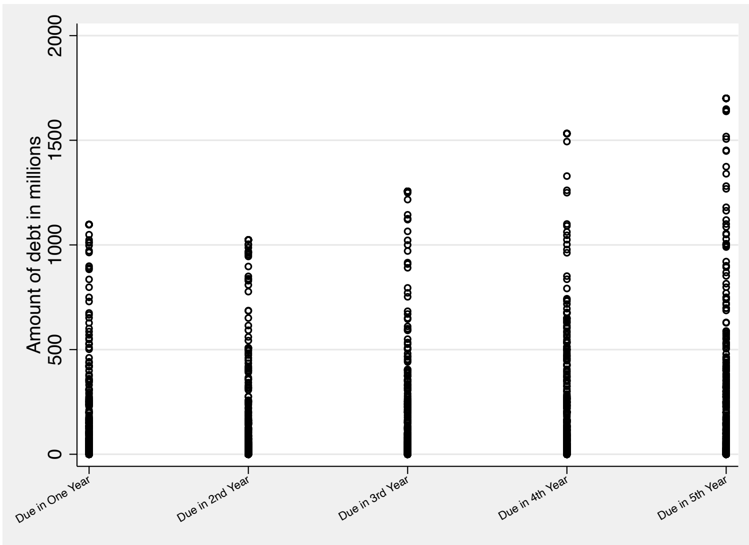
Figure OA B1: Counties on the states' borders



Notes: Figure OA B1 shows counties located on the states' borders. The areas shaded in dark green represent pairs of counties that share a border and have their centroids within a 75 km distance of each other.

Corporate debt structure during the financial crisis. Figure OA B2 illustrates the wide variation in long-term debt maturity across different years at the beginning of the financial crisis.

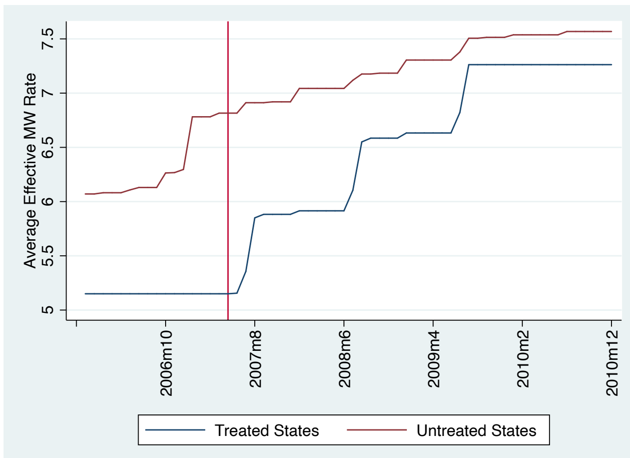
Figure OA B2: Corporate debt structure



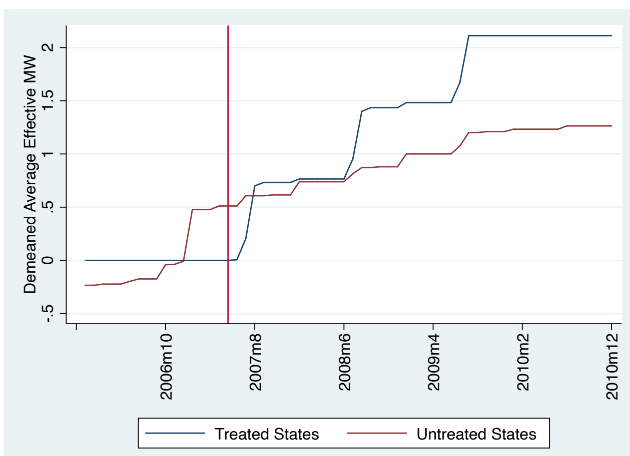
Notes: Figure OA B2 illustrates the wide variation in debt maturity across different years at the beginning of the financial crisis.

Effective minimum wage changes of treated and untreated states. We show in Figure OA B3(a) the effective average minimum wage in states treated by the new federal minimum wage in 2007 and states untreated by the new federal minimum wage. Figure OA B3(b) shows the divergence patterns for the two groups when we demean each time series using the pre-crisis average for the treatment and the control groups.

Figure OA B3: Effective minimum wage changes of treated and untreated states



(a) Effective MW rates

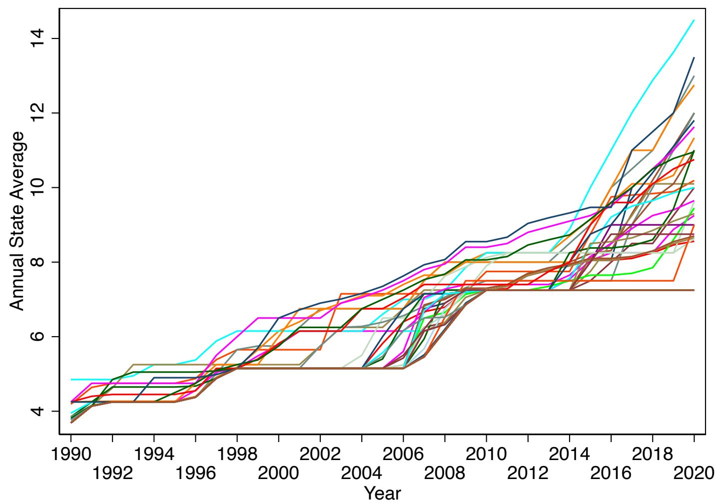


(b) Divergence in the effective MW rates

Notes: Figure OA B3(a) shows the effective average minimum wage in states treated by the increase in the federal minimum wage in 2007 and states untreated by the new federal minimum wage rate. Figure OA B3(b) shows the divergence patterns for the two groups when we demean each time series using the pre-crisis mean for the treatment and control groups.

Minimum wage rate across states. Figure OA B4 shows the staggered changes in minimum wage states across states from 1990 to 2020 for each state. Our empirical analysis takes advantage of this significant variability in minimum wage changes across states and periods.

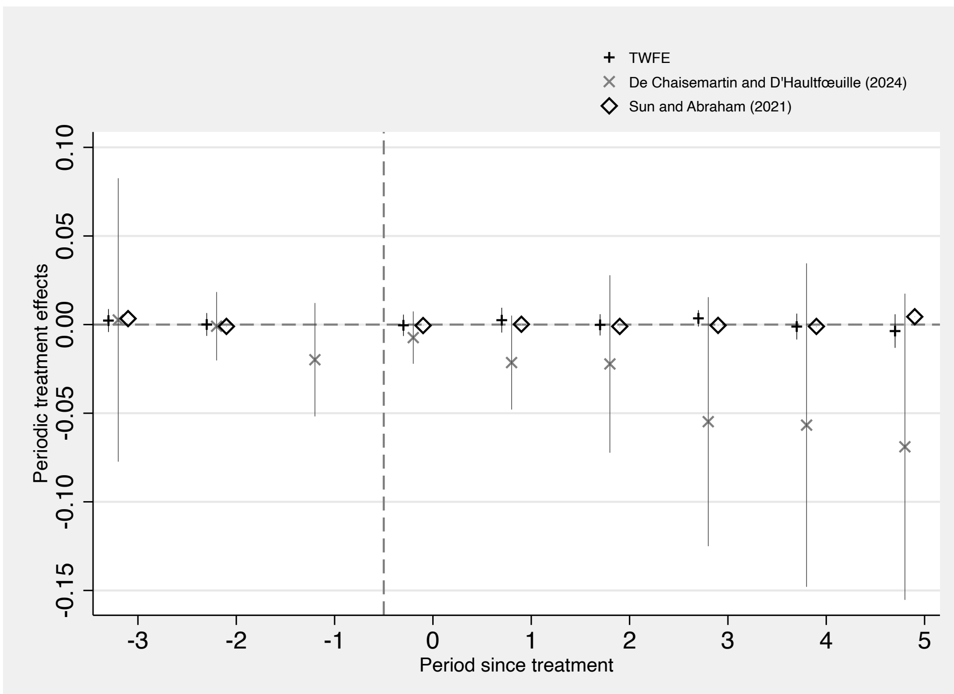
Figure OA B4: Minimum wage dynamics across states



Notes: Figure OA B4 shows changes in minimum wage policies for each state during our period of analysis.

Alternative staggered difference-in-differences estimators. We focus our analysis on counties located along states' borders and present in Figure OA B5 annual point estimates of the effect of minimum wage changes on employment obtained using the methodologies introduced by De Chaisemartin and d'Haultfoeuille (2020) and Sun and Abraham (2021). We also compare these estimators to the results of a two-way fixed effects (TWFE) model.

Figure OA B5: Alternative estimators



Notes: Figure OA B5 depicts the annual treatment effects utilizing the two-way fixed effects (TWFE) estimator and compares them with the estimators introduced by De Chaisemartin and d'Haultfoeuille (2020) and Sun and Abraham (2021). The outcome variable is the natural logarithm of employment in the establishment. We focus on counties along the states' borders. The plot exhibits yearly point estimates and 95% confidence intervals based on standard errors clustered by state.

Online Appendix B Tables

Summary statistics. Our final database comprises 2,340,503 establishment-year observations, 231,552 establishments, and 5,615 firms. We analyze these data over the period spanning 1990 to 2020. We report summary statistics for this sample in Panel A of Table OA B1. We also aggregated the data to the firm level to further evaluate our findings and gain insight into how exposure to minimum wage adjustments impacts firm performance and policies. This database comprises 54,829 firm-year observations spanning the period 1990-2020. We report the summary statistics in Panel B of Table OA B1.

Table OA B1: Descriptive statistics

	(1) Count	(2) Mean	(3) SD	(4) p25	(5) p50	(6) p75
Panel A: Establishment-level database						
Employment (log)	2,340,503	2.8332	1.4740	1.7918	2.7081	3.8286
MW	2,340,503	6.3455	1.7173	5.1500	6.1579	7.2500
Log(Firm Size)	2,314,911	7.6621	2.1169	6.1942	7.6285	9.0865
WW Index	2,254,576	-0.3697	0.1121	-0.4472	-0.3716	-0.2940
SA index	2,314,911	-0.9798	1.4520	-2.2112	-1.3329	-0.0687
ROA	2,282,050	0.0462	0.1043	0.0145	0.0569	0.0980
ROI	2,279,778	0.0969	5.7201	0.0215	0.0900	0.1602
Tangibility	2,279,165	0.3921	0.2428	0.1932	0.3749	0.5457
Employment growth	2,081,402	0.3019	0.8002	-0.0667	0.1351	0.3804
Cash volatility	2,286,569	0.3727	0.7640	0.0883	0.1755	0.3324
Distress	2,339,787	0.1099	0.3127	0.0000	0.0000	0.0000
Cash holdings	2,238,807	0.0861	0.1015	0.0192	0.0471	0.1168
Employment beta	2,332,840	2.5008	6.4046	0.3685	0.9563	2.4106
Constraining Words (log)	1,512,346	5.7267	0.7479	5.2832	5.7366	6.1944
KZ Index	1,809,761	-0.5511	2.7808	-0.8537	0.0993	0.8171
Log(Sales)	2,340,491	14.6562	1.6451	13.4794	14.3502	15.8047
Lerner Index	2,174,144	0.1228	0.1023	0.0601	0.0999	0.1775
Panel B: Firm-level database						
Employment (log)	54,829	1.2178	1.1803	0.2429	0.8381	1.8825
Exposure MW	54,829	6.0921	1.6927	4.8904	5.4465	7.3073
Log(Firm Size)	54,829	5.5462	2.2023	3.9859	5.5431	7.0551
SA Index	54,829	0.7236	2.0432	-0.8510	0.5255	2.0423
WW Index	53,452	-0.2591	0.1255	-0.3431	-0.2613	-0.1806
Liquidity	48,941	0.2016	0.2874	0.0257	0.0926	0.2596
Leverage (Long)	48,827	0.2342	0.2862	0.0045	0.1494	0.3492
Leverage (Short)	48,951	0.0625	0.1443	0.0002	0.0122	0.0537
R&D	29,296	0.0943	0.1580	0.0000	0.0313	0.1196
Capital Expenditures	48,572	0.0626	0.0772	0.0175	0.0374	0.0759
ROA	48,942	-0.0489	0.3122	-0.0545	0.0304	0.0806

Notes: This table shows descriptive statistics of the variables used in our analysis. See Table OA B2 for a detailed description of our variables and their sources.

Variable definition. Table OA B2 contains detailed information on the variables that we use in the empirical analysis, their description, and their sources.

Table OA B2: Variable description

Variable name	Description	Source
Panel A: Variables used in the establishment-level analysis		
Log(Employment)	The natural logarithm of the number of employees working at the establishment.	NETS
Treated	An indicator variable equal to one if the establishment is located in a bounded state to the federal minimum wage change, zero otherwise.	Authors
Constraints (short)	Ratio of firm's long-term debt due in one year (item DD1) to sales (item SALE). This variable is measured in the fiscal year 2007. We report the value in % by multiplying it by 100.	Compustat
Cash holdings	Ratio of cash (item CH) to lagged total assets (item AT).	Compustat
ROA	Ratio of firm's income before extraordinary items (item IB) to lagged total assets (item AT).	Compustat
ROI	Ratio of the firm's net income (loss) (item NI) to lagged invested capital (item ICAPT).	Compustat
Tangibility	Ratio of a firm's property, plant, and equipment (item PPENT) to lagged total assets (item AT).	Compustat
Historical short-term debt	Average level of the firm's short-term debt across all the sample years, measured as a ratio of long-term debt due in 1 year (item DD1) to net sales/turnover (item SALE), multiplied by 100.	Compustat
Employment growth	Employment growth rate over the last four years.	Compustat
Cash flow volatility	The standard deviation of cash flows divided by the absolute value of the mean cash flows over the previous four years.	Compustat
Distress	Indicator variable equal to 1 if its earnings before interest, taxes, depreciation, and amortization (EBITDA) are lower than its interest expenses for two consecutive years, or if its EBITDA is less than 80 percent of its interest expenses in any single year.	Compustat
Constraints (long)	Ratio of firm's long-term debt due in two, three, four, and five years (item DD2, DD3, DD4, and DD5) to sales (item SALE). This variable is measured as in the fiscal year 2007. We report the value in % by multiplying it by 100.	Compustat

Notes: This table shows a detailed description of the variables and their data sources.

Table OA B2: Variable description (continued)

MW	State-level effective minimum wage.	Vaghul and Zipperer (2021)
Log(Firm Size)	Natural logarithm of one plus the firm's total assets in millions of dollars (item AT).	Compustat
SA Index	Following Hadlock and Pierce (2010), the SA index is defined as (-0.737 $\times$ log(AT))+(0.043 $\times$ log(AT)^2)-(0.040 $\times$ Age). Age is defined as the number of years the firm is present in Compustat. Size is winsorized at (the log of) \$4.5 billion, and Age is winsorized at 37 years. A higher value of the index indicates greater financial constraints.	Compustat
WW Index	Following Whited and Wu (2006), the WW index is defined as (-0.091 $\times$ CF) - (0.062 $\times$ DIVPOS)+(0.021 $\times$ TLTD) - (0.044 $\times$ log(AT)) + (0.102 $\times$ ISG) - (0.035 $\times$ SG), where CF is the ratio of cash flow (item IB + item DP) to total assets (item AT), DIVPOS is an indicator variable equal to one if the firm pays cash dividends (item DVT) and zero otherwise, TLTD is the ratio of long-term debt (item DLTT) to total assets (item AT), ISG is the firm's 3-digit SIC industry sales (item SALE) growth, SG is firm sales (item SALE) growth. A higher value of the index indicates greater financial constraints.	Compustat
KZ Index	Following Kaplan and Zingales (1997), the KZ index is defined as: -1.001$\times$[(IB+DP)/PPEGT] +0.282$\times$[(AT+MktVal-CEQ-TXDB)/AT]+3.139 $\times$[(DLTT+DLC)/(DLTT+DLC+AT)]-39.367$\times$[(DVC+DVP)/PPEGT]-1.314$\times$[CH/PPEGT]. MktVal is defined as [CSHO$\times$PRCC]. MktVal and PPEGT are lagged. A higher value of the index indicates greater financial constraints.	Compustat
Composite Index	A combined measure of corporate financial constraints defined as equal to one if the majority of the firm's financial constraint indicators are equal to one (above-median lagged annual values for the KZ, SA, and WW indices; below-median annual values for lagged firm size), and zero otherwise.	Compustat
PCA Index	An index based on the principal component analysis (PCA) of the four measures of financial constraints (KZ, SA, WW indices, and Size).	Compustat
Constraining Words (log)	A textual analysis measure of financial constraint based on the dictionary of words associated with	

Notes: This table describes the variables and their data sources.

Table OA B2: Variable description (continued)

	constraint in 10-K. The measure is defined as the natural logarithm of one plus the number of constraining words in the firm's 10-K.	Bodnaruk et al. (2015)
Bond Constraint	An indicator variable equal to one if the firm did not issue a bond in a given year, and zero otherwise.	FISD
SL Constraint	An indicator variable equal to one if the firm did not obtain a syndicated loan (SL) in a given year, and zero otherwise.	DealScan
Population (log)	Natural logarithm of a county's population as reported by the Bureau of Economic Analysis (BEA).	BEA
Income Per Capita	County's income per capita as reported by the Bureau of Economic Analysis (BEA).	BEA
Earnings Per Capita	County's earnings per capita as reported by the Bureau of Economic Analysis (BEA).	BEA
Employment Ratio	County's employment-population ratio as reported by the Bureau of Economic Analysis (BEA).	BEA
Sales	Total establishment's sales in dollar amounts.	NETS
Lerner Index	A measure of pricing power of a firm reflecting price-cost margin (PCM) calculated as (SALE-COGS-XSGA)/SALE.	Compustat
	Panel B: Variables used in the corporate-level analysis	
Employment (log)	Natural logarithm of one plus total firm employees, measured in thousands (item EMP).	Compustat
Treated	Share of employees affected by the federal minimum wage increase.	NETS
MW Exposure	Corporate exposure to minimum wage policies are measured as state minimum wages weighted by the number of employees at the firm's establishments across the states.	Vaghul and Zipperer (2021)
Log(Firm Size)	Natural logarithm of one plus the firm's total assets in millions of dollars (item AT).	Compustat
SA Index	Following Hadlock and Pierce (2010), the SA index is defined as (-0.737 $\times$ log(AT))+(0.043 $\times$ log(AT)^2)-(0.040 $\times$ Age).	

Notes: This table describes the variables and their data sources.

Table OA B2: Variable description (continued)

	Age is defined as the years the firm is present in Compustat. Size is winsorized at (the log of) \$4.5 billion, and Age is winsorized at 37 years. A higher value of the index indicates a greater financial constraint.	Compustat
WW Index	Following Whited and Wu (2006), the WW index is defined as (-0.091 $\times$ CF) - (0.062 $\times$ DIVPOS)+(0.021 $\times$ TLTD) - (0.044 $\times$ log(AT)) + (0.102 $\times$ ISG) - (0.035 $\times$ SG), where CF is the ratio of cash flow (item IB + item DP) to total assets (item AT), DIVPOS is an indicator variable equal to one if the firm pays cash dividends (item DVT) and zero otherwise, TLTD is the ratio of long-term debt (item DLTT) to total assets (item AT), ISG is the firm's 3-digit SIC industry sales (item SALE) growth, SG is firm sales (item SALE) growth. A higher value of the index indicates a greater financial constraint.	Compustat
Liquidity	Ratio of cash and and short-term investments (item CHE) to lagged total assets (item AT).	Compustat
Leverage (Short)	Ratio of debt in current liabilities (item DLC) to lagged total assets (item AT).	Compustat
Leverage (Long)	Ratio of long-term debt (item DLTT) to lagged total assets (item AT).	Compustat
R&D	Ratio of research and development expenses (item XRD) to lagged total assets (item AT).	Compustat
Capital Expenditures	Ratio of capital expenditures (item CAPX) to lagged total assets (item AT).	Compustat
ROA	Return on assets ratio measured as earnings before extraordinary items (item IB) divided by lagged total assets (item AT).	Compustat

Notes: This table describes the variables and their data sources.

The financial crisis sample. We report the summary statistics for the financial crisis sample in Table OA B3. We limit the spanning period of analysis to the period 2003-2011.

Table OA B3: Descriptive statistics - Financial crisis sample

	(1)	(2)	(3)	(4)	(5)	(6)
	Count	Mean	SD	p25	p50	p75
Employment (log)	714,091	2.8056	1.4698	1.6094	2.7081	3.8067
Treated	714,091	0.3436	0.4749	0.0000	0.0000	1.0000
Constraints (short) (%)	611,659	1.1986	2.2373	0.0089	0.2995	1.4141
Constraints (long) (%)	553,591	10.6749	15.6378	0.3172	3.4949	13.7329

Notes: This table shows descriptive statistics of the variables used in our financial crisis analysis. We limit the sample period to 2003-2011. See Table OA B2 for a detailed description of the variables and their data sources.

Employment dynamics after the federal minimum wage increase. We estimate the impact of the federal minimum wage on employment using Equation (2) and report the estimates for the full sample, the sample of counties along state borders, and county pairs located within 75 km of each other in Table OA B4.

Table OA B4: Difference-in-differences - Financial crisis

	(1) All counties	(2) Counties on borders	(3) Counties on borders ($\leq$ 75 km)
	Employment (log)	Employment (log)	Employment (log)
Post × Treated	-0.006	-0.015***	-0.014**
	(0.004)	(0.005)	(0.006)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	714,091	217,077	192,622
R-squared	0.962	0.964	0.965

Notes: This table shows regression results for Equation (2). We use *Employment (log)* as our outcome variable. All the regressions include establishment and year × firm fixed effects. We focus on three alternative samples: (i) all counties in the United States in Column (1), (ii) all counties on state borders in Column (2), and (iii) all counties on state borders whose centroids are less than 75 km apart in Column (3). Robust standard errors, clustered at the state level, are reported in parentheses below the coefficient estimates. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of the variables and their data sources.

Additional results. As shown in Column (1) of Table OA B5, our results remain robust when we include the historical short-term debt as a control and interact it with both the post-period and minimum wage treatment indicators. In Column (2), we further demonstrate that the results continue to hold when we include and fully interact all other corporate characteristics with the treatment indicators. In Columns (3) and (4), we show that measurement errors are unlikely to affect our main findings. To further test the validity of our approach, we decompose the Post variable into individual years using a generalized difference-in-differences setting to analyze the employment dynamics around the Great Recession crisis period for financially constrained firms. More specifically, we report each yearly interaction coefficient in Column (5) of Table OA B5. As a falsification test, we investigate in column (6) the impact of long-term debt maturing after the end of the financial crisis, denoted as *Constraints (long)*.

Table OA B5: Additional results

	(1)	(2)	(3)	(4)	(5)	(6)
	Employment (log)	Employment (log)	Employment (log)	Employment (log)	Employment (log)	Employment (log)
Post × Treated	-0.008	-0.024	-0.019	-0.033	-0.025	-0.015*
	(0.007)	(0.021)	(0.017)	(0.020)	(0.021)	(0.007)
Post × Constraints (short)	-0.006**	-0.008**	-0.005*	-0.007**		
	(0.003)	(0.003)	(0.003)	(0.003)		
Post × Constraints (long)						0.000
						(0.000)
Year 2004 × Treated × Constraints (short)					-0.001	
					(0.003)	
Year 2005 × Treated × Constraints (short)					-0.000	
					(0.004)	
Year 2006 × Treated × Constraints (short)					-0.007	
					(0.004)	
Year 2007 × Treated × Constraints (short)					-0.005	
					(0.004)	
Year 2008 × Treated × Constraints (short)					-0.010**	
					(0.004)	
Year 2009 × Treated × Constraints (short)					-0.011***	
					(0.004)	
Year 2010 × Treated × Constraints (short)					-0.014***	
					(0.005)	
Year 2011 × Treated × Constraints (short)					-0.014***	
					(0.005)	
Establishment FE	✓	✓	✓	✓	✓	✓
Firm × Year FE	✓	✓	✓	✓	✓	✓
Fully interacted historical short-term debt	✓	✓	✓	✓	✓	
Other fully interacted controls		✓	✓	✓	✓	
Observations	185,947	167,971	145,372	140,188	167,971	168,518
R-squared	0.962	0.963	0.975	0.965	0.963	0.965

Notes: This table reports robustness analysis and additional empirical tests examining the relationship between federal minimum wage increases and firm-level employment, focusing on counties on borders sample. Column (1) controls for the historical use of short-term debt. Column (2) includes full interactions between treatment indicators and several firm characteristics (total assets, return on assets, return on investment, asset tangibility, growth rate of employment over the past four years, cash-flow ratio, cash flow volatility, and the distress indicator). In Columns (3) and (4), we show that measurement errors are unlikely to affect our main findings. Column (5) presents a generalized difference-in-differences specification with year-specific interactions. Column (6) reports on a falsification test using long-term debt maturing after the Great Recession (*Constraints (long)*). Robust standard errors, reported in parentheses, are clustered at the state level. *p ¡ 0.10, **p ¡ 0.05, ***p ¡ 0.01. See Table OA B2 for a detailed description of the variables and their data sources.

Weights used in the counterfactual analysis. The size categories of the establishments we consider in our analysis are grouped in intervals as follows: less than 1-9 employees, 10-19 employees, 20-29 employees, 30-39 employees, 40-49 employees, up to more than 100 employees. To calculate the weights, we consider the employment shares of these establishments in 2006 from the entire NETS database. We present the weights we utilize in Table OA B6.

Table OA B6: Weights

Employment size category	Weights
1-9 Employees	0.25
10-19 Employees	0.09
20-29 Employees	0.06
30-39 Employees	0.05
40-49 Employees	0.04
50-59 Employees	0.04
60-69 Employees	0.03
70-79 Employees	0.02
80-89 Employees	0.02
90-99 Employees	0.01
More than 100	0.38

Notes: The table shows the weights per employment category that we use for our counterfactual analysis.

Employment and minimum wage policies. To evaluate the impact of a minimum wage hike on employment, we estimate Equation (8) in the main text, excluding the interaction term. In our baseline specification, we restrict our sample to counties located on state borders. Additionally, we report results for all counties in the United States, as well as for county pairs in different states that share a border and have centroids within 75 km of each other. We report the results in Table OA B7.

Table OA B7: Difference-in-differences - Average effect

	(1) All counties Employment (log)	(2) Counties on borders Employment (log)	(3) Counties on borders ($\leq$ 75 km) Employment (log)
MW	-0.002	0.001	-0.003
	(0.002)	(0.003)	(0.004)
Establishment FE	✓	✓	✓
Year $\times$ Firm FE	✓	✓	✓
Observations	2,340,503	707,713	625,679
R-squared	0.936	0.939	0.939

Notes: This table shows regression estimates for Equation (8) in the main text, excluding the interaction term. We use *Employment (log)* as our outcome variable. We focus on three alternative samples: (i) all counties in the United States in Column (1), (ii) all counties on state borders in Column (2), and (iii) all counties on state borders whose centroids are less than 75 km apart in Column (3). Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of our variables and their sources.

Revenues as corporate financing mechanism. Using establishment-level sales data, we find in Table OA B8 no average increase in revenues following minimum wage hikes. When we account for firm financial constraints, the results indicate that constrained firms experience a decline in sales, consistent with these firms cutting employment and output rather than raising prices. This pattern echoes our broader findings that financially constrained firms face limited scope to pass through cost shocks and instead rely more heavily on internal cash preservation and external financing. In contrast, we document revenue gains only for firms with greater pricing power, as measured by the firm-level Lerner index, suggesting that price pass-through is feasible only in specific competitive environments and normal economic times.

Table OA B8: Minimum wages - Sales

	(1)	(2)	(3)	(4)	(5)
	Sales (log)	Sales (log)	Sales (log)	Sales (log)	Sales (log)
MW	0.001	-0.026***	-0.022**	-0.008**	-0.005
	(0.002)	(0.009)	(0.009)	(0.003)	(0.003)
MW × Log(Firm Size)		0.003***			
		(0.001)			
MW × WW Index			-0.056***		
			(0.021)		
MW × SA Index				-0.006***	
				(0.002)	
MW × Lerner Index					0.030**
					(0.014)
Establishment FE	✓	✓	✓	✓	✓
Year × Firm FE	✓	✓	✓	✓	✓
Observations	707,706	690,714	661,772	690,714	654,569
R-squared	0.950	0.950	0.951	0.950	0.952

Notes: This table shows regression results for Equation (8) for counties on borders sample. We use *Sales (log)* as our outcome variable. In the last column, we interact the minimum wage increase with the Lerner index measured at the firm level. All the regressions include the establishment and year × firm fixed effects. We focus on all counties on state borders. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of the variables and their data sources.

Other corporate characteristics. We present evidence that our primary results hold when we incorporate additional corporate controls in our regression analysis: firm performance, efficiency, asset tangibility, the cash-to-asset ratio, a distress indicator, and the corporate growth rate of employment over the past four years. In addition, we control for the employment beta, defined as a statistic that quantifies the relationship between the time series variability of employment within a four-digit SIC code and the variability of total employment. We report our results in Table OA B9.

Table OA B9: Confounding corporate characteristics

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
MW	-0.027**	-0.025**	-0.007*
	(0.011)	(0.010)	(0.004)
MW × Log(Firm Size)	0.003**		
	(0.001)		
MW × WW Index		-0.059**	
		(0.026)	
MW × SA Index			-0.004*
			(0.002)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Fully interacted corporate controls	✓	✓	✓
Observations	611,376	596,657	611,376
R-squared	0.940	0.941	0.940

Notes: This table shows regression results for Equation (8) for counties on borders sample. We use *Employment (log)* as our outcome variable. All the regressions include the establishment and year × firm fixed effects. We also control for several major corporate characteristics (total assets, return on assets, return on investment, asset tangibility, firm's growth rate in employment over the past four years, cash flow ratio, and industry employment sensitivity to shocks, measured by a firm's employment beta). Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of the variables and their data sources.

County characteristics. We demonstrate that our results remain robust even when explicitly controlling for several local economic characteristics that we document in Table OA B10 to be influenced by minimum wage policies. More specifically, we control for a county's population, employment ratio, earnings per capita, and income per capita, and report our results in Table OA B11.

Table OA B10: Minimum wage policies and changes in county characteristics

	(1)	(2)	(3)	(4)
	Population (log)	Employment Ratio	Income Per Capita	Earnings Per Capita
MW	0.003	0.003	678.253**	416.711*
	(0.008)	(0.002)	(276.048)	(234.931)
County FE	✓	✓	✓	✓
Year FE	✓	✓	✓	✓
Observations	34,596	34,594	34,596	34,596
R-squared	0.996	0.792	0.920	0.869

Notes: This table shows the results of a regression of minimum wage rate on alternative county characteristics for the counties in the border sample. All regressions include county and year fixed effects. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *$p<0.10$, **$p<0.05$, ***$p<0.01$. See Table OA B2 for a detailed description of the variables and their data sources.

Table OA B11: Controlling for county characteristics

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
MW	-0.028***	-0.024***	-0.008**
	(0.010)	(0.009)	(0.004)
MW × Log(Firm Size)	0.003***		
	(0.001)		
MW × WW Index		-0.063***	
		(0.020)	
MW × SA Index			-0.006***
			(0.002)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
County controls	✓	✓	✓
Observations	677,761	649,451	677,761
R-squared	0.939	0.940	0.939

Notes: This table shows regression results for Equation (8) for counties on borders sample. We use *Employment (log)* as our outcome variable. All the regressions include the establishment and year × firm fixed effects. We also control for a county's earnings per capita and income per capita. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of the variables and their data sources.

Controlling for other state policies. We gather data on the unemployment insurance (UI) benefit schedules of each state from the U.S. Department of Labor's "Significant Provisions of State UI Laws". More specifically, we follow Guo et al. (2024) and consider the overall UI benefit level in a given state and year by multiplying the maximum weekly benefit amount by the maximum duration of benefits provided under each state's regular UI program. Additionally, we collect information on union membership rates (UMR) from the Current Population Survey (CPS), and corporate tax changes, median income tax rate, and top income tax rate from Baker et al. (2025). We include these controls in Equation (8). The outcomes are reported in Table OA B12.

Table OA B12: Minimum wage and other state policies

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
MW	-0.026***	-0.024***	-0.009**
	(0.010)	(0.009)	(0.004)
MW × Log(Firm Size)	0.003***		
	(0.001)		
MW × WW Index		-0.062***	
		(0.021)	
MW × SA Index			-0.006***
			(0.002)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Controlling for other state policies	✓	✓	✓
Observations	669,860	641,661	669,860
R-squared	0.940	0.941	0.940

Notes: This table shows regression results for Equation (8) for counties on borders sample. We use *Employment (log)* as our outcome variable. All the regressions include establishment and year × firm fixed effects. We control for alternative state-level policies, including unemployment insurance policies, union membership rates, corporate tax changes, median income tax rate, and top income tax rate. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of the variables and their data sources.

Matching of paired counties. We follow Dube et al. (2010) and propose an alternative identification strategy. We match each county on a state border to potentially multiple pairs of other counties. We then control for pair-by-period fixed effects. The results are reported in Table OA B13. To further test the robustness of our results, we introduce a more advanced approach recently proposed by Jha et al. (2025). More specifically, they demonstrate that a positive bias arises in a county-pair specification when pairs are formed using contiguous counties across state lines, but where the pairs are in different commuting zones (CZ). We report these results in Table OA B14.

Table OA B13: County pairing

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
MW	-0.026**	-0.020**	-0.010**
	(0.010)	(0.010)	(0.005)
MW × Log(Firm Size)	0.003**		
	(0.001)		
MW × WW Index		-0.044**	
		(0.021)	
MW × SA Index			-0.005***
			(0.002)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Pair Counties × Year FE	✓	✓	✓
Observations	1,614,208	1,547,462	1,614,208
R-squared	0.943	0.943	0.943

Notes: This table shows regression results for Equation (8). We use *Employment (log)* as our outcome variable. All the regressions include the establishment and year × firm fixed effects. Furthermore, we alternatively include pair counties fixed effects and pair counties × year fixed effects. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of our variables and their data sources.

Table OA B14: Commuting-zone pairing

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
MW	-0.025*	-0.023**	-0.004
	(0.014)	(0.011)	(0.006)
MW × Log(Firm Size)	0.003**		
	(0.001)		
MW × WW Index		-0.063***	
		(0.023)	
MW × SA Index			-0.004
			(0.003)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
CZ-state × Year FE	✓	✓	✓
Observations	703,892	673,739	703,892
R-squared	0.941	0.942	0.941

Notes: This table shows regression results for Equation (8). We use *Employment (log)* as our outcome variable. All the regressions include the establishment and year × firm fixed effects. Furthermore, we alternatively include CZ-state times year fixed effects. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of our variables and their data sources.

Measurement errors in the NETS database. We conduct a robustness check by removing establishments with fewer than three employees, as these observations are more likely to have imputed data. In an additional test, we remove establishments with round numbers of employees (5, 10, 100, 200, ... 1000), as these observations are also more likely to be based on imputation. The results are reported in Tables OA B15 and OA B16.

Table OA B15: Removing establishments with less than three employees

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
MW	-0.025***	-0.020***	-0.006*
	(0.008)	(0.007)	(0.003)
MW × Log(Firm Size)	0.003***		
	(0.001)		
MW × WW Index		-0.055***	
		(0.018)	
MW × SA Index			-0.005***
			(0.002)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	633,733	607,055	633,733
R-squared	0.953	0.954	0.953

Notes: This table shows regression results for Equation (8) for counties on borders sample. We use *Employment (log)* as our outcome variable. All the regressions include establishment and firm time year fixed effects. Additionally, we remove establishments with fewer than three employees. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of the variables and their data sources.

Table OA B16: Removing establishments with round numbers of employees

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
MW	-0.027***	-0.023***	-0.007*
	(0.009)	(0.008)	(0.003)
MW × Log(Firm Size)	0.003***		
	(0.001)		
MW × WW Index		-0.060***	
		(0.021)	
MW × SA Index			-0.005**
			(0.002)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	575,366	550,687	575,366
R-squared	0.942	0.943	0.942

Notes: This table shows regression results for Equation (8) for counties on borders sample. We use *Employment (log)* as our outcome variable. All the regressions include establishment and firm time year fixed effects. Additionally, we remove establishments that report a round number of employees (5, 10, 100, 200, ... 1000), as these are more likely to involve imputation. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of the variables and their data sources.

Alternative measures of financial constraint. We investigate whether our results hold when considering alternative measures of financial constraints. We start by considering a commonly used financial constraint measure in the corporate finance literature, that is, the KZ index (Kaplan and Zingales, 1997). We report the results in Column (1) of OA B17. To further assess the robustness of our findings, we adopt a methodology similar to that presented by Bartram et al. (2022) by constructing a composite financial constraints measure. This method relies on the rankings of firms based on the KZ, SA, and WW indices and their asset size. We categorize a firm as financially constrained if it exhibits above-median annual values for the KZ, SA, and WW indices alongside below-median annual values for firm size. If most of these indicators point to financial constraints, the composite indicator is set to one; otherwise, it is set to zero. We present the results in Column (2) of Table OA B17. Our results still hold when we consider a composite index based on the principal component analysis (PCA) of these four alternative measures of financial friction. We report the results in Column (3) of Table OA B17. Finally, we demonstrate the robustness of our results when considering an alternative measure of financial constraint based on a textual analysis of a firm's annual 10-K filing. We use the dictionary of 'constraining' words proposed in Bodnaruk et al. (2015) to do so. We take the natural logarithm of one plus the word frequency of this dictionary and control for differences in the length of the 10-K filings. We present the results in Column (4) of Table OA B17. Finally, in the last two columns, we use whether a firm issued a bond or obtained a syndicated loan (SL) as a proxy for financial constraints.

Table OA B17: Financial constraints - Alternative variables

	(1)	(2)	(3)	(4)	(5)	(6)
	Emp. (log)	Emp. (log)	Emp. (log)	Emp. (log)	Emp. (log)	Emp. (log)
MW	0.002	0.006*	-0.001	-0.107	0.003	0.006*
	(0.003)	(0.003)	(0.003)	(0.070)	(0.003)	(0.003)
MW × KZ Index	-0.000					
	(0.001)					
MW × Composite Index		-0.009**				
		(0.004)				
MW × PCA Index			-0.005***			
			(0.002)			
MW × Constraining Words(log)				-0.012*		
				(0.006)		
MW × SL Constraint					-0.003**	
					(0.002)	
MW × Bond Constraint						-0.007***
						(0.002)
Establishment FE	✓	✓	✓	✓	✓	✓
Year × Firm FE	✓	✓	✓	✓	✓	✓
Observations	544,259	543,217	543,217	453,328	707,713	707,713
R-squared	0.942	0.942	0.942	0.952	0.939	0.939

Notes: This table shows regression results for Equation (8) for counties on borders sample. We use *Employment (log)* as our outcome variable and alternative measures of financial constraints as our independent variables. All the regressions include establishment and year × firm fixed effects. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. **p* <0.10, ***p* <0.05, ****p* <0.01. See Table OA B2 for a detailed description of our variables and their sources.

Potential factors influencing the positive effects of minimum wage changes on employment. We use alternative metrics for local labor market conditions to more precisely test whether and when establishments of unconstrained firms increase employment following a minimum wage rise. We explore the heterogeneous effects on employment based on the potential supply of minimum wage workers in Table OA B18, hiring gap in Table OA B19, labor turnover in Table OA B20, and corporate labor market power in Table OA B21.

Table OA B18: Potential supply of minimum wage workers

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
		Panel A: Below median labor pool	
MW	-0.000	0.001	0.002
	(0.005)	(0.004)	(0.004)
MW × Unconstrained - Log(Firm Size)	0.005		
	(0.006)		
MW × Unconstrained - WW Index		0.004	
		(0.005)	
MW × Unconstrained - SA Index			0.002
			(0.004)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	364,732	349,236	364,732
R-squared	0.942	0.943	0.942
		Panel B: Above median labor pool	
MW	-0.009**	-0.008	-0.013**
	(0.004)	(0.005)	(0.005)
MW × Unconstrained - Log(Firm Size)	0.015***		
	(0.004)		
MW × Unconstrained - WW Index		0.014*	
		(0.007)	
MW × Unconstrained - SA Index			0.023***
			(0.005)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	314,197	301,758	314,197
R-squared	0.941	0.942	0.941

Notes: This table shows the estimation results from Equation (9) for the two subsamples created by splitting counties located on state borders at the median value of the potential supply of minimum wage workers. We use *Employment (log)* as our outcome variable. All the regressions include establishment and year × firm fixed effects. Robust standard errors, reported in parentheses, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of the variables and their data sources.

Table OA B19: Hiring gap

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
		Panel A: Below the median hiring gap	
MW	-0.002	-0.003	-0.004
	(0.005)	(0.004)	(0.006)
MW × Unconstrained - Log(Firm Size)	-0.001		
	(0.007)		
MW × Unconstrained - WW Index		0.001	
		(0.008)	
MW × Unconstrained - SA Index			0.002
			(0.010)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	296,189	285,175	296,189
R-squared	0.943	0.944	0.943
		Panel B: Above the median hiring gap	
MW	-0.004	-0.002	-0.003
	(0.004)	(0.004)	(0.004)
MW × Unconstrained - Log(Firm Size)	0.012**		
	(0.006)		
MW × Unconstrained - WW Index		0.010*	
		(0.006)	
MW × Unconstrained - SA Index			0.011**
			(0.005)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	385,029	367,952	385,029
R-squared	0.940	0.940	0.940

Notes: This table shows the estimation results from Equation (9) for the two subsamples created by splitting counties located on state borders at the hiring gap median value. We use *Employment (log)* as our outcome variable. All the regressions include establishment and year × firm fixed effects. Robust standard errors, reported in parentheses, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01.

Table OA B20: Labor turnover

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
		Panel A: Below median labor turnover	
MW	-0.000	0.003	0.004
	(0.006)	(0.004)	(0.005)
MW × Unconstrained - Log(Firm Size)	0.006		
	(0.006)		
MW × Unconstrained - WW Index		0.002	
		(0.004)	
MW × Unconstrained - SA Index			-0.002
			(0.006)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	296,094	284,248	296,094
R-squared	0.943	0.943	0.943
		Panel B: Above median labor turnover	
MW	-0.012**	-0.010*	-0.012**
	(0.005)	(0.005)	(0.005)
MW × Unconstrained - Log(Firm Size)	0.017***		
	(0.005)		
MW × Unconstrained - WW Index		0.015**	
		(0.006)	
MW × Unconstrained - SA Index			0.018***
			(0.005)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	382,553	366,595	382,553
R-squared	0.940	0.941	0.940

Notes: This table shows the estimation results from Equation (9) for the two subsamples created by splitting those counties located on state borders at the median value of labor turnover. We use *Employment (log)* as our outcome variable. All the regressions include establishment and year × firm fixed effects. Robust standard errors, reported in parentheses, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01.

Table OA B21: Labor market concentration

	(1) Employment (log)	(2) Employment (log)	(3) Employment (log)
		Panel A: Below median HHI	
MW	-0.000	0.001	-0.001
	(0.004)	(0.004)	(0.003)
MW × Unconstrained - Log(Firm Size)	0.011**		
	(0.005)		
MW × Unconstrained - WW Index		0.011**	
		(0.005)	
MW × Unconstrained - SA Index			0.011***
			(0.004)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	304,444	291,728	304,444
R-squared	0.936	0.937	0.936
		Panel B: Above median HHI	
MW	-0.004	-0.004	-0.001
	(0.005)	(0.005)	(0.004)
MW × Unconstrained - Log(Firm Size)	0.005		
	(0.006)		
MW × Unconstrained - WW Index		0.004	
		(0.006)	
MW × Unconstrained - SA Index			-0.000
			(0.006)
Establishment FE	✓	✓	✓
Year × Firm FE	✓	✓	✓
Observations	373,760	359,003	373,760
R-squared	0.943	0.944	0.943

Notes: This table shows the estimation results from Equation (9) for the two subsamples created by splitting establishments in counties on state borders at the median value of corporate labor market power (as measured by HHI). We use *Employment (log)* as our outcome variable. All the regressions include establishment and year × firm fixed effects. Robust standard errors, reported in parentheses, are clustered at the state level. $*p$ <0.10, $**p$ <0.05, $***p$ <0.01.

County level employment analysis. To account for general equilibrium effects in our empirical analysis, we adopt the methodology outlined by Giroud and Mueller (2017) by employing county-level regressions, as reported in Equation (10). The findings in Table OA B22 echo our baseline results.

Table OA B22: County level employment regressions

	(1)	(2)	(3)
	Employment (log)	Employment (log)	Employment (log)
MW	-0.290***	-0.225***	-0.011
	(0.063)	(0.059)	(0.031)
Log(Firm Size)	-0.134***		
	(0.045)		
MW × Log(Firm Size)	0.036***		
	(0.006)		
WW Index		2.294***	
		(0.711)	
MW × WW Index		-0.595***	
		(0.116)	
SA Index			0.160**
			(0.066)
MW × SA Index			-0.040***
			(0.011)
County FE	✓	✓	✓
Year FE	✓	✓	✓
Observations	28,323	27,874	28,323
R-squared	0.892	0.895	0.890

Notes: This table shows regression results for Equation (10) for counties on borders sample. We use *Employment (log)* as our outcome variable. All the regressions include the county and year fixed effects. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of our variables and their sources.

Establishments of private firms. We use the entire NETS database, which also contains information on private establishments and the headquarter to which they belong. We estimate the following specification to be consistent with the estimation of our main set of results.

$$Log(Employment)_{i,t} = \beta_0 MW_{s,t} + \beta_1 MW_{s,t} \times Private_{f,t} + \delta_{f,t} + \eta_i + \epsilon_{i,t} \quad \text{(OA10)}$$

In this setting, *Private* is an indicator variable equal to one if an establishment belongs to a private firm and zero otherwise. β_0 is the impact of minimum wage policies on employment at publicly owned establishments, while β_1 is the heterogeneous impact of minimum wages on employment at privately owned establishments compared to public firm establishments. We present the estimation results in Table OA B23 for the alternative samples considered in our empirical analysis. We observe that β_0 is negative but not statistically significant. Importantly, consistent with our hypothesis, β_1 is negative and statistically significant when we focus on the sample of establishments located on county borders.

Table OA B23: Establishments of private firms

	(1) All Counties	(2) Counties on borders	(3) Counties on borders ($\leq$ 75 km)
	Employment (log)	Employment (log)	Employment (log)
MW	-0.000	0.002	0.001
	(0.001)	(0.002)	(0.002)
MW $\times$ Private	-0.001	-0.004***	-0.004***
	(0.001)	(0.001)	(0.001)
Establishment FE	Yes	Yes	Yes
Year $\times$ Firm FE	Yes	Yes	Yes
Observations	68,975,411	20,714,884	18,340,120
R-squared	0.922	0.924	0.924

Notes: This table shows regression results for Equation (OA10). We use *Employment (log)* as our outcome variable. The financial constrained measure is an indicator equal to one if the establishment belong to a private firm. All the regressions include the establishment and year $\times$ firm fixed effects. We focus on three alternative samples: (i) all counties in the United States in Column (1), (ii) all counties on state borders in Column (2), and (iii) all counties on state borders whose centroids are less than 75 km apart in Column (3). Robust standard errors, clustered at the state level, are reported in parentheses below the coefficient estimates. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of our variables and their data sources.

Employment measurement issues in Compustat. We exclude multinational firms from the sample. We classified a firm as multinational if it reported non-zero foreign income in the previous three years. Alternatively, we defined a firm as multinational if at least 5% of its sales were from outside the home country (according to the Compustat Geographic Segment database) (Nimier-David et al., 2023). We report in these tables results consistent with our primary evidence in the main text.

Table OA B24: Firm level analysis of employment - Excluding multinational firms based on the foreign income

	(1) Employment (log)	(2) Employment (log)	(3) Employment (log)
MW Exposure	-0.030**	-0.016	0.004
	(0.014)	(0.013)	(0.015)
Log(Firm Size)	0.196***		
	(0.016)		
MW Exposure × Log(Firm Size)	0.006**		
	(0.003)		
WW Index		-0.705***	
		(0.177)	
MW Exposure × WW Index		-0.078**	
		(0.033)	
SA Index			-0.125***
			(0.017)
MW Exposure × SA Index			-0.007**
			(0.003)
Firm FE	✓	✓	✓
Year × Industry FE	✓	✓	✓
Observations	34,570	30,840	34,570
R-squared	0.962	0.955	0.957

Notes: This table shows regression results for Equation (12). We use *Employment (log)* as our outcome variable. All the regressions include firm and year × industry fixed effects. We exclude from the sample firms that report non-zero foreign income in the previous three years. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. **p* <0.10, ***p* <0.05, ****p* <0.01. See Table OA B2 for a detailed description of our variables and their data sources.

Table OA B25: Firm level analysis of employment - Excluding multinational firms based on sales from foreign countries

	(1) Employment (log)	(2) Employment (log)	(3) Employment (log)
MW Exposure	-0.068***	-0.042**	-0.001
	(0.017)	(0.020)	(0.018)
Log(Firm Size)	0.203***		
	(0.016)		
MW Exposure × Log(Firm Size)	0.010***		
	(0.002)		
WW Index		-0.721***	
		(0.179)	
MW Exposure × WW Index		-0.124***	
		(0.038)	
SA Index			-0.107***
			(0.017)
MW Exposure × SA Index			-0.017***
			(0.003)
Firm FE	✓	✓	✓
Year × Industry FE	✓	✓	✓
Observations	25,039	22,439	25,039
R-squared	0.966	0.959	0.961

Notes: This table shows regression results for Equation (12). We use *Employment (log)* as our outcome variable. All the regressions include firm and year × industry fixed effects. We exclude from the sample firms that report at least 5% of their sales from outside the home country. Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of our variables and their data sources.

Minimum wage and the cost of labor. We gather wage data from the Current Business Patterns (CBP). This database offers total salary information at the county, year, and industry levels.[23] With this data, we estimate the following Equation:

$$Wage_{j,k,t} = \alpha_j + \theta_{k,t} + \beta MW_{s,t} + \epsilon_{j,k,t} \tag{OA11}$$

The dependent variable of interest is the total payroll divided by the number of employees in the county. We also control for county (α_j) and industry-year fixed effects ($\theta_{k,t}$). We report the estimates for the alternative samples in Table OA B26.

Our findings reveal that, on average, minimum wage increases lead to a rise in the cost of labor per employee at the county and year levels. Specifically, we observe an increase of about $900 per employee per year. This result aligns well with prior research indicating a positive impact of minimum wage policies on average wages. More specifically, minimum wage policies affect not only minimum wage workers, but they also have spillover effects on other relatively low-wage incumbent workers and new hires (e.g., Gopalan et al., 2021, 2024).

[23] After 1998, NAICS codes are used to report data at the industry level. Before 1998, industry information was reported using SIC codes. Therefore, we convert the four-digit SIC code to a six-digit NAICS code using the 1997 bridge file.

Table OA B26: Wages

	(1) All counties	(2) Counties on borders	(3) Counties on borders ($\leq$ 75 km)
	Wage	Wage	Wage
MW	704.152***	764.570***	929.820***
	(157.282)	(209.026)	(322.189)
County FE	✓	✓	✓
Year $\times$ Industry FE	✓	✓	✓
Observations	7,374,781	2,521,090	2,203,509
R-squared	0.366	0.459	0.467

Notes: This table shows regression results for Equation (OA11). We use *Wage* as our outcome variable. We focus on three alternative samples: (i) all counties in the United States in Column (1), (ii) all counties on state borders in Column (2), and (iii) all counties on state borders whose centroids are less than 75 km apart in Column (3). Robust standard errors, reported in parentheses below the coefficient estimates, are clustered at the state level. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for a detailed description of our variables and their sources.

Cash holdings, long-term debt, and constrained firms. We explicitly test whether unconstrained firms are the ones using liquidity and long-term debt to finance higher wage costs. Specifically, we use firm-level cash holdings and long-term debt as the outcome variable and interact minimum-wage exposure with terciles of our financial-constraint measures. We report the results in Table OA B27.

Table OA B27: Corporate policies by financial constraints

	(1)	(2)	(3)
Financial Constraint Variable	Log(Firm Size)	WW Index	SA Index
Panel A: Outcome - Liquidity			
MW | Constraint Variable = 0	0.003	-0.008**	-0.013***
	(0.005)	(0.004)	(0.004)
2nd tercile Financial Constraint × Mean MW	-0.011***	-0.005*	0.006**
	(0.003)	(0.002)	(0.003)
3rd tercile Financial Constraint × Mean MW	-0.013***	0.003	0.021***
	(0.004)	(0.003)	(0.005)
Firm FE	✓	✓	✓
Industry × Year FE	✓	✓	✓
Observations	48,422	47,286	48,422
R-squared	0.611	0.605	0.610
Panel B: Outcome - Long-term Debt			
MW | Constraint Variable = 0	0.007	0.012**	0.013**
	(0.005)	(0.005)	(0.005)
2nd tercile Financial Constraint × Mean MW	0.001	-0.005	-0.006
	(0.005)	(0.004)	(0.004)
3rd tercile Financial Constraint × Mean MW	0.008	-0.002	-0.005
	(0.006)	(0.005)	(0.007)
Firm FE	✓	✓	✓
Industry × Year FE	✓	✓	✓
Observations	48,308	47,237	48,308
R-squared	0.595	0.601	0.595

Notes: This table reports regression results for Equation (13). Panel A uses *Cash Holdings* as the dependent variable, while Panel B uses *Long-term Debt.* In addition, we interact the minimum wage rate with the terciles of the distribution of the alternative financial-constraint variables. All regressions include firm and year × industry fixed effects. Robust standard errors, clustered at the state level, are reported in parentheses below coefficient estimates. *p <0.10, **p <0.05, ***p <0.01. See Table OA B2 for detailed variable definitions and data sources.

Online Appendix C Matching establishment, minimum wage, and corporate financial friction information

The 2020 version of the National Establishments Time Series (NETS) provides the legal business names of establishments, along with the names of the primary domestic firms within a hierarchical "Family Tree" of companies. To ensure the precise inclusion of relevant balance sheet attributes from the individual firms themselves, rather than their parent holding companies, we employ the legal business names of these establishments for cross-referencing with the Compustat database in our empirical analysis.

We utilize fuzzy matching, employing a similar threshold of 90%, to align company names between the two databases. Subsequently, we conduct manual verification to ensure the precision of these matches. With this methodology, we successfully merged 353,818 establishments to obtain a dataset of 4,231,721 establishment-year observations, each containing no missing information on employment levels or geographical locations.

We integrate this database with historical minimum wage data sourced from Vaghul and Zipperer (2021), leading to the exclusion of 6,348 establishment-year records. Among these are public corporation establishments located in Puerto Rico or the Virgin Islands, for which minimum wage information is unavailable. Subsequently, we merged the database with corporate balance sheet attributes sourced from Compustat, resulting in a dataset comprising 2,819,629 observations.

To refine the dataset, we remove establishments of the firms from the financial and utility sectors (SIC codes 60 and 49), eliminating 312,309 and 95,526 observations, respectively. Furthermore, we employ the *reghdfe* command from Correia (2016) to eliminate singleton observations, reducing the sample by an additional 62,291 observations.

Our final database is composed of 2,340,503 establishment-year observations, 231,552 establishments, and 5,615 companies from 1990 to 2020.

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